Modelling the effect of common mental disorders on child growth in Butajira, Ethiopia

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By

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Statement of Authorship

The research carried out in this thesis was part of a larger study funded as a research training fellowship for Dr Charlotte Hanlon (Wellcome Trust Clinical Tropical Medicine). I spent most of the data collection period in Butajira, Ethiopia, where the project was conducted. Charlotte and I design the study questionnaire and trained the data collectors. We jointly participated in supervision of data collection, data entry and data cleaning on a day-to-day basis. For the results presented in this thesis, I developed the analysis plan with close guidance from Prof Michael Dewey, conducted the analyses and am responsible for interpretation of findings.

Abstract

Background: Evidence about the effects of perinatal common mental disorders (CMD) on child growth is consistent in South Asian studies but not in Sub Saharan Africa.

Aims: To (1) assess the effect of CMD on child growth in Ethiopia using traditional analysis, latent growth modelling (LGM) and multilevel growth modelling (MGM) techniques, and (2) evaluate the effects of other pre-specified risk factors on infant growth using these three modelling techniques

Methods: A population based cohort was established between July 2005 and February 2006 in a demographic surveillance site in Butajira, Ethiopia, recruiting 1065 women in pregnancy and followed them with their newly born infants. Main exposure was perinatal CMD measured with locally validated self reporting questionnaire (SRQ-20). The women were interviewed at recruitment, birth, two and 12 months postnatal. In addition to birthweight, infant growth was monitored at two, six, nine, 12 and 18 months of age and standardized using 2006 WHO reference standards. Logistic regression and linear regression were used to model binary and continous infant growth outcomes, respectively, at two, six and twelve months of age. Furthermore, infant growth over the first 18 months of age and predictors of growth patters of these infants were investigated using MGM and LGM.

Results: Postnatal and persistent CMD were significant risk factors of compromised initial infant length, and part of the effect on the total length gain was mediated through diarrhoeal episodes. Early infant feeding practices and birthweight did not mediate the effects of CMD on infant growth. Boys perform better in initial weight (in kg) and length (in cm), but worse in initial values of standardized growth measures. Low birthweight and reduced maternal mid upper arm circumference during pregnancy were significant predictors of compromised initial growth. These effects and the effect of gender were maintained during the follow-up period. Rural residence and higher maternal autonomy score were significant predictors of better initial length and the latter was positively associated with the total gain in length during infancy. Higher score on poor sanitary condition scale was significantly associated with better initial length and compromised initial weight. However, it was inversely associated with the change in weight and length during infancy. Of the three modelling techniques, LGM has more flexibility and produces easily interpretable results.

Conclusions: In our study setting perinatal CMD has significant negative effect on infant growth. Other risk factors of infant growth are composed of infant, maternal and environmental characteristics. LGM should be the first choice to model complex relationships like these.

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Abbreviations

AIC:	Akaike Information criteria
AMOS:	Analysis of moment structures
BIC:	Bayesian Information criteria
BRHP:	Butajira Rural Health Programme
CFA:	Confirmatory factor analysis
CFI:	Comparative fit index
CHA:	Community health agents
CIDI:	Composite International Diagnostic Interview
CI:	Confidence interval
CMD:	Common mental disorders
DSS:	Demographic surveillance site
EFA:	Exploratory factor Analysis
FIML:	Full information maximum likelihood
FMOH:	Federal Ministry of Health
GEE:	Generalized estimating equations
GLM:	Generalized linear models
GLS:	Generalized least squares
HEP:	Health extension package
HEW:	Health extension workers
HMIS:	Health management information system
HSDP:	Health sector development program
IEC:	Information, education and communication
LAMIC:	Low and middle income countries
LGM:	Latent growth modelling
LISREL:	Linear structural relations
LMP:	Last menstrual period
MAR:	Missing at random
MCAR:	Missing completely at random
MGM:	Multilevel growth modelling
ML:	Maximum likelihood
MUAC:	Mid upper arm circumference
NFI:	Normed fit index

NNFI:	Non-normed fit index
OLS:	Ordinary least squares
PA:	Path analysis
PCA:	Principal component analysis
P-MaMiE:	Perinatal Maternal Mental Disorder in Ethiopia
SEM:	Structural equation modelling
SD:	Standard deviation
SNNPR:	Southern Nations, Nationalities and People's Region
SRA:	Structural regression analysis
SRMEA:	Standardized root mean square error approximation
SRQ-20:	Self-reporting questionnaire, 20-item version
SSA:	Sub-Saharan Africa
TBA:	Traditional birth attendant
TLI:	Tucker-Lewis index
UNDP:	United nations development program
UNICEF:	United nationals children's fund
WHO:	World health organization

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CHAPTER 1: BACKGROUND

1.1 Introduction

The main objective of this chapter is to give a background of the study to understand why the study was conducted. The details are organized in the following order:

- Description of the main outcome, undernutrition , and highlights of commonly used anthropometric measures which are also adopted in this study
- Description of the global public health importance of undernutrition
- Description of the magnitude and trends of undernutrition in Ethiopia
- Review of risk factors for child undernutrition
- Definition of perinatal common mental disorder (CMD) and their association with child undernutrition
- Current practice in modelling the effect of CMD on child growth including a critical review of current practices
- Alternative modelling techniques to assess the effect of CMD on child growth

1.2. Undernutrition and anthropometric measures

1.2.1 Undernutrition

Undernutrition refers to a deficiency of nutrition (Blo¨ssner and de Onis 2005) which creates an imbalance in the components of the body composition required for growth in comparison to a healthy person of similar age and sex (WHO Expert Committee 1995). The reasons behind these deficits include but are not limited to inadequate food intake, inadequate composition of required components in daily food, poor absorption of the required nutrients, increased excretion from the body, and repeated infections (Beisel 1977; WHO Working Group 1986; Blo¨ssner and de Onis 2005; Black, Allen et al. 2008). Unfavourable consequences of undernutrition vary from compromising quality of life to becoming an underlying cause of death (Gómez, Galvan et al. 2000; Ezzati, Lopez et al. 2002; Blo¨ssner and de Onis 2005; Black, Allen et al. 2008; Victora, Adair et al. 2008). To address this critical public health problem in a meaningful way requires a combination of overall economic growth at all levels, identification of locally relevant risk factors and targeted programmatic interventions (Haddad, Alderman et al. 2003; Bhutta, Ahmed et al. 2008)

1.2.1.1 Anthropometry as a measure of undernutrition

The importance of having cheap and easy-to-implement standardized measures of nutritional status has long been promoted by the World Health Organization (WHO) (Jelliffe 1966). In line with this several reference standards have been used at various times although their applicability to the population in developing countries was debated (Garn 1962; Garn 1965; Habicht, Yarbrough et al. 1974). These reference standards are mainly based on anthropometric measurements which are easy to measure. The frequent and wider use of anthropometric measurements to define child health and nutritional status is because of the sensitivity of child growth to any disturbances in health and nutrition regardless of their aetiology (Blo¨ssner and de Onis 2005).

Anthropometric measurements have been an important parameter for paediatricians in evaluating the overall health and well-being of children at an individual level and for public health officers to monitor growth and overall health at a group level (WHO Expert Committee 1995). A deficit in weight was in use for a long time as a measure to identify and target treatment of undernourished infants (Gómez 2000) until the publication of new international weight and height-based growth reference standards in the mid 1970s which had the objective of distinguishing between acute and chronic, or present and past, undernutrition (Waterlo, Buzina et al. 1977). The motivation to have one international growth reference standard was based on the similarity of growth patterns among well-fed and healthy pre-school children with regards to weight as well as length/height, at least for the first six months of age, regardless of ethnic differences in different continents (Habicht, Yarbrough et al. 1974). The WHO has been revising the growth reference standards (WHO Working group 1986; WHO Expert Committee 1995) and the most recent growth reference standard was published in 2006 (WHO 2006)

1.2.1.2 Anthropometric measures, indices and indicators

The WHO definition distinguishes these three related terms (WHO Expert Committee 1995). Anthropometric measures are numerical measurements that can be obtained by direct measurement. Length/height, weight, mid-upper arm circumference, and head circumference are examples of anthropometric measures. Individual anthropometric

measures are believed to be less useful than combinations of the measures, as will now be discussed (Garn 1962; WHO Working Group 1986; WHO Expert Committee 1995)

Anthropometric indices are mathematical combinations of anthropometric measures that have been used as indicators of nutritional status. Length/height-for-age, weight-for-age and height-for-weight are the most commonly used anthropometric indices to assess child nutritional status. These indices can be expressed in terms of z-scores (or standard deviation score), percentiles or percent of median (WHO Expert Committee 1995). Of the three indices length/height-for-age is a measure of chronic undernutrition, weight-for-height is a measure of acute undernutrition and weight-for-age does not distinguish between the two types of undernutrition. This shows the usefulness of the first two measures in cross-sectional surveys and weight-for-age in case of serial measurements (Waterlo, Buzina et al. 1977)

For any one of the three indices a z-score can be generated by subtracting the median growth measure of the reference population corresponding to the age of the measured child from the observed growth measure of the child, and dividing that difference by the standard deviation of the reference population of the same age. To get a percentile for a given child the observed growth measure is evaluated against the growth measure of the reference population of the same age. This gives the percent of the reference population with the same or less than the growth measure of the target child which is equal to the percentile of the child. The third index, percent of median, is generated by obtaining the percentage of the ratio of the observed anthropometric measure to the median of the reference population of the same age. Because of its mathematical properties the current recommendation is to use the z-score (WHO 2006).

The term indicator is related to the use or application of the indices in a given area of interest (WHO Expert Committee 1995). For example, the proportion of children who are more than two standard deviation below the mean weight-for-height z score indicates the level of wasting at that particular time and could be considered to be an indicator of current food shortage in that area. Indicators are useful for governments, non-governmental international organizations and clinicians for practical application which includes risk assessment, targeted intervention, evaluation of intervention, evaluation of level of development, individual based specialized treatment (WHO 2006). For example,

"underweight" (weight-for-age z score < two standard deviations below the median) is one of the indicators used in estimating global burden of diseases (Ezzati, Lopez et al. 2002) and in the assessment of success of the millennium development goals (United Nations 2001). The recommendations accompanying the new 2006 WHO growth reference standard did not make any change to the previous practice of using -2SD and -3SD as the cut-off value of z-scores to demarcate mild and severe undernutrition, respectively (WHO 2006).

1.2.2 Undernutrition as a global public health problem

Child undernutrition is highly prevalent (Black, Allen et al. 2008) especially in low and middle income countries (LAMIC) and is a serious public health problem as a result of its association with high morbidity and mortality (Blo¨ssner and De Onis 2005; Victora, Adair et al. 2008). Childhood and maternal underweight ranks first in the global burden of diseases (Ezzati, Lopez et al. 2002). In order to tackle the magnitude and the consequences of these and other global public health problems there is an ongoing worldwide effort focused on the complete eradication of extreme poverty and hunger (UNICEF. 2002) with strong worldwide political commitment (United Nations 2001). However, the burden of undernutrition is still a major public health problem especially in resource poor countries (Muller and Krawinkel 2005; Black, Allen et al. 2008). Ninety percent of the world's stunted children live in 36 developing countries (Bhutta, Ahmed et al. 2008; Black, Allen et al. 2008). Undernutrition remains a major cause of disability and mortality (World Bank 2006), ranked as the top cause of global burden of disease (Ezzati, Lopez et al. 2002) and underlying 53% of deaths in children under five years (Bryce, Boschi-Pinto et al. 2005; Muller and Krawinkel 2005). The potential negative impact of child undernutrition goes beyond the individual, affecting society and future generations (Ramakrishnan, Martorell et al. 1999; Grantham-McGregor, Cheung et al. 2007; Victora, Adair et al. 2008).

Despite an encouraging global downward trend in the prevalence of stunting, the progress is not uniform across countries (de Onis 2000; Muller and Krawinkel 2005) According to global estimates taking the most recent available data for the years 2000-2006, the prevalence of underweight and stunting among under five children in sub-

Saharan Africa were 28% and 38% respectively, while among least developed countries in general it was 35% and 42% respectively (UNICEF 2008).

1.3 Undernutrition in Ethiopia

Ethiopian history has documented repeated episodes of drought and famine, each time claiming significant number of lives and exposing the majority of survivors to an acute shortage of resources although the public health impact of these repeated incidents are not fully documented (Taye, Haile Mariam et al. 2010). A non-systematic review of the undernutrition literature in Ethiopia showed that there was an increasing trend of child undernutrition between 1983 and 1998, with notable variability between regions (Getahun, Urga et al. 2001). However, since 1998, four consecutive countrywide surveys have shown a decreasing trend in the prevalence of undernutrition, both underweight and stunting, in all age and sex categories (Central Statistical Authority 2004). A different analysis of countrywide surveys also showed a decreasing trend of underweight undernutrition between 1996 and 2005, but the prevalence remained high even by Horn of Africa standards (Mason, Chotard et al. 2010). More recent estimates show this downward trend continuing (Central Statistical Agency and ORC Macro 2006). Nevertheless, the prevalence of underweight (38%) and stunting (47%) among under five children remains substantial (UNICEF 2008).

1.4 Risk factors for child undernutrition

Epidemiological studies conducted in developing countries have identified several factors associated with undernutrition, including low parental education, parental undernutrition, poverty and food insecurity, low maternal intelligence, lack of maternal autonomy within the family, rural residential area, unhygienic household environment, parasitic infection, sub-optimal infant feeding practices, low birth weight, increasing child age, and child being male (Getahun, Urga et al. 2001; Muller and Krawinkel 2005; Silva 2005; Wachs, Creed-Kanashiro et al. 2005; Kumar, Goel et al. 2006; Casapı´a, Joseph et al. 2007; Huntsman and White 2007; Wamani, Åstrøm et al. 2007; Bhutta, Ahmed et al. 2008; Engebretsen, Tylleskar et al. 2008; Wachs 2008; Bomela 2009; Shroff, Griffiths et al. 2009; Subramanian, Ackerson et al. 2009). Maternal depression or psychological morbidity (Prince, Patel et al. 2007; Surkan, Kennedy et al.

2011) are also emerging as important determinants of child undernutrition, particularly in studies from South Asia.

Analysis of data from three consecutive welfare monitoring surveys in Ethiopia over the period 1996-1998 (Christiaensen and Alderman 2001) identified low household resources, lower parental education, high food prices and low maternal nutritional knowledge as key determinants of growth faltering in children. A study focused on the southern region of Ethiopia (Yimer 2000), which included the current study site, identified low socioeconomic status of household, low maternal education, short previous birth interval, having many children aged under five years and older age of infant as risk factors for child undernutrition. The possible reasons for the high prevalence of undernutrition in the age group of 12 to 23 months was hypothesised to be a combination of insufficient and inappropriate supplementary foods, and recurrent infections such as diarrhoea due to poor sanitation (Getahun, Urga et al. 2001). In a one year follow-up of a large birth cohort in western Ethiopia, good sanitary conditions, availability of a water supply, increased family income and maternal literacy were associated with weight gain, while the traditional surgical practices of uvulectomy and milk teeth extraction were both associated with reduction of infant weight (Asefa, Hewison et al. 1998). Male sex has also been identified as one of the risk factors for undernutrition among Ethiopian children (Central Statistical Authority 2004) (Central Statistical Agency and ORC Macro 2006).

1.5 Perinatal common mental disorders (CMD)

1.5.1 Defining Perinatal CMD

Biological, psychological and social changes that women face during pregnancy, child birth and the postnatal period may predispose them to emotional distress, even though the expectation of having a child may be a source of happiness for the rest of the family. Research into perinatal maternal mental disorders has been predominantly focused on mental health problems of the postnatal period (Lee and Chung 2007; Mann, Gilbody et al. 2010). Pregnancy is now also recognised as a time when women may experience a multiplicity of mental disorders (Andersson, Sundström-Poromaa et al. 2003) although most studies have concentrated on depressive disorders (Bennett, Einarson et al. 2004), psychosocial stress or pregnancy-related anxiety (Paarlberg, Vingerhoets et al. 1999)

Early research in the field highlighted the distinctive presentation of non-psychotic depression occurring in the postnatal period (Pitt 1968). However, subsequent work has found postnatal depression to be indistinguishable from depression occurring at any other time in a woman's life (Cox, Connor et al. 1982), and the degree to which it is possible to separate out distinct depressive, anxiety or somatoform disorders in ante- and postnatal women has been questioned (Stewart, Kauye et al. 2009). In non-perinatal women from community and primary care settings, there are high levels of co-variance of symptoms of depressive, anxiety and somatoform disorders, indicating an underlying unitary construct to the CMD (Lewis, Pelosi et al. 1992; Goldberg 1996).

In perinatal women, the focus of studies is typically narrowed to ante- or postnatal depression, although evidence supports the relevance of a broader concept of perinatal CMD (Aderibigbe, Gureje et al. 1993; Matthey, Barnett et al. 2003). Perinatal CMD is highly prevalent in developed countries (Bennett, Einarson et al. 2004; Gavin, Gaynes et al. 2005) and developing countries (Cooper, Tomlinson et al. 1999; Patel, Rodrigues et al. 2002; Rahman, Iqbal et al. 2003; Abiodun 2006; Husain, Bevc et al. 2006; Mohammad, Gamble et al. 2011). The negative effect of perinatal CMD could severely affect the wellbeing of the mother and the newborn, and it could be a hurdle for societal development (Prince, Patel et al. 2007).

1.5.2 SRQ-20 for measurement of perinatal CMD

The 20 item version of the self reporting questionnaire (SRQ-20) was originally developed by the World Health Organization (WHO) to measure CMD in primary healthcare settings (Harding, de Arango et al. 1980). The 20 items include depressive, anxiety and somatic symptoms. SRQ-20 has been widely used across LAMICs (Beusenberg and Orley 1994) and recommended as a cost-effective indicator of community mental health (Harpham, Reichenheim et al. 2003). Hence, SRQ-20 could be considered as an alternative research tool in settings where mental health professionals are in short supply and not readily available to participate in research projects (Alem and Kebede 2003).

The SRQ-20 has been validated in several LAMICs for detection of perinatal CMD (Rahman, Iqbal et al. 2005; Stewart, Umar et al. 2008; Weobong, Akpalu et al. 2009) and in women of child-bearing age (Aidoo and Harpham 2001; Pollock, Manaseki-Holland et al. 2006). The recommended optimal SRQ-20 cut-off score for detecting perinatal CMD ranges from ≥5 to ≥9 in validation studies with sensitivity ranging between 77% and 95% and specificity ranging between 63% and 84%). In a review of SRQ validation studies for detection of non-perinatal CMD the range of recommended cut-off scores ranges from ≥4 to ≥12 and the authors note that different cut-offs are required for different study settings (Beusenberg and Orley 1994)

No published studies from Ethiopia have specifically measured CMD in the perinatal period. With the exception of a small subgroup of community sample included in one validation study (Kortmann 1990) all other validation studies of SRQ in Ethiopia are based on small sample size recruited from health institution clinics (Kortmann and ten Horn 1988; Kortmann 1990; Zilber, Youngmann et al. 2004; Youngmann, Zilber et al. 2008). However, SRQ-20 has been used in several large community surveys in Ethiopia with different cut-off scores used to indicate presence of CMD; ≥6 (Kebede, Alem et al. 1999; Mogga, Prince et al. 2006), ≥8 (Tadesse, Kebede et al. 1999; Harpham, Huttly et al. 2005) and ≥11 (Tafari, Aboud et al. 1991; Mulatu 1995; Alem, Kebede et al. 1999). The scale has been found to be an acceptable and feasible method for quantifying the burden of CMD in Ethiopia at the community level. Hence, SRQ-20 is a good candidate to measure perinatal CMD in rural Ethiopia after validating and establishing appropriate cut-off scores to indicate cases of perinatal CMD (Beusenberg and Orley 1994).

1.6 Perinatal CMD and the newborn

1.6.1 The effect of antenatal CMD on Birth Weight

Existing evidence indicates that antenatal CMD may be a risk factor for low birth weight (Steer, Scholl et al. 1992; Copper, Goldenberg et al. 1996; Sheehan 1998; Paarlberg, Vingerhoets et al. 1999; Field, Diego et al. 2004; Patel, Rahman et al. 2004; Rahman, Bunn et al. 2007; Grote, Bridge et al. 2010). However, significant negative effect of antenatal CMD on birth weight was not replicated in other studies (Collin, Chisenga et al.

2006; Hanlon, Medhin et al. 2009). The mechanisms through which antenatal CMD may affect neonatal outcomes in general and birthweight in particular are likely to be multifactorial and indirect (McAnarney and Stevens-Simon 1990; Copper, Goldenberg et al. 1996; Rahman, Harrington et al. 2002). Exposure to antenatal CMD might influence maternal behaviour choices and some of these choices could potentially affect the fetus (Rahman, Harrington et al. 2002). Although more evidence is needed to fully understand the role of maternal biochemistry and hormonal activities as mediators of the effect of CMD on pregnancy outcomes (Steer, Scholl et al. 1992), elevated norepinephrine level was reported as a significant mediator with regard to low birthweight (Field, Diego et al. 2004).

The relative risk of low birthweight associated with antenatal CMD is estimated to be 2.05 in developing nations, 1.16 in United States and 1.15 at a global level (Grote, Bridge et al. 2010). Comorbidity of antenatal depression with anxiety worsened the negative effect on prematurity but the incidence of low birthweight remained unaffected (Field, Diego et al. 2010). Since low birth weight is an established risk factor for impaired child development and mortality (Shenkin, Starr et al. 2001; Lawn, Cousens et al. 2005), the reviewed evidence implies an indirect effect of antenatal CMD on other birth outcomes. In high-income countries, maternal depression during pregnancy is associated with poor attendance at antenatal clinics, low birth weight and preterm delivery (Pagel, Smilkstein et al. 1990; Hedegaard M , Henriksen TB et al. 1993; Copper, Goldenberg et al. 1996). In low-income countries where access to antenatal care is often difficult, maternal depression could influence the level of care received, indirectly increasing the incidence of low birth weight, and subsequent infant mortality and morbidity (Rahman, Iqbal et al. 2003; Patel, Rahman et al. 2004).

Birth weight is a strong predictor of early childhood weight and height, the association being relatively stronger with childhood weight (Binkin, Yip et al. 1988). In several studies conducted in LAMIC, low birth weight has been shown to be associated with subsequent infant undernutrition (de Miranda, Turecki et al. 1996; Saleemi, Ashraf et al. 2001; Anoop, Saravanan et al. 2004; Santos, Matijasevich et al. 2010). In Brazil, 27% of infants with Protein Energy undernutrition, compared to 6% of the controls infants had been of low birth weight (de Miranda, Turecki et al. 1996). In another population-based cohort in Brazil, low birth weight and preterm birth were significantly associated with

increased prevalence of underweight, stunting, and wasting at four years of age (Santos, Matijasevich et al. 2010). In India, undernourished children were 2.9 times more likely to have had a low birth weight than the control group of normal weight infants (Anoop, Saravanan et al. 2004). In a nested case-control study in Pakistan, maternal height, birth-weight and stunting at six months of age were significant predictors of stunting at 12 months of age (Saleemi, Ashraf et al. 2001). Synthesis of the available research evidence at a global level supports the link between low birth weight and undernutrition, and their individual and joint contribution to infant and child mortality (Black, Morris et al. 2003; Lawn, Cousens et al. 2005). Based on the reviewed evidence related to the three characteristics (maternal depression, low birth weight and infant growth), significant direct and indirect effect of maternal depression on child growth can be hypothesised.

1.6.2 The effect of antenatal CMD on early infant feeding practices

The current recommendation for the well-being of the newborns includes immediate initiation of beast feeding, avoidance of any prelacteal feeding, providing colostrum to the newborn and exclusive breast feeding for the first 4 to 6 months (WHO 1989). If followed, these practices have the potential to reduce infant morbidity and enhance growth (Islam, Ahmed et al. 2006; Diallo, Bell et al. 2009). However, adherence to these recommendations is likely to be influenced by the characteristics of individual parents, different values of the community, coverage of health service delivery infrastructure and level of economic development of the country. In developed counties the environment is supportive for the implementation of these recommendations because the health delivery system is well established and the coverage of health professional-attended delivery is reasonably high. However, in developing countries where the health delivery system is generally below standard, not easily accessible because of poor infrastructure and the coverage of health professional attended delivery is still below 50% (WHO 2010), implementation of optimal infant feeding recommendations is challenging.

In developing countries, cultural beliefs and practices of individual parents and the community at large are more likely to have a significant influence on early infant feeding practices. A study conducted among a displaced population in the southern Turkey (Ergenekon-Ozelci, Elmaci et al. 2006) found that (a) mothers had a generally positive attitude towards breastfeeding, (b) mothers with lower education generally believed that

colostrum should not be fed to the infant because a pregnant woman's milk is unhealthy for the baby, (c) no woman was found to feed her infant exclusively by breastfeeding, (d) only 9.9% of mothers initiated breastfeeding within the first hour of birth, and (e) 40.0% of mothers started solid foods before four months. A qualitative study conducted in the highlands of the northern Lao Peoples Democratic Republic reported several cultural beliefs operating as barriers to exclusive breastfeeding, including food taboos and hygiene behaviours which could potentially explain the high prevalence of undernutrition in the area (Holmes, Damian Hoy et al. 2007). In a cross-sectional survey in Turkey, including 921 woman whose youngest child was 6 to 18 months old, nearly all mothers had breastfed their infants at some time, 62.2% delayed initiation of breastfeeding for at least 24 hours, and almost half of the infants received sweetened water before initiation of breast feeding (Saka, Ertem et al. 2005).

There is some evidence in the literature emerging from cross-sectional studies showing that deviation from recommended early feeding practices might be linked with infant undernutrition. In Turkey (Ergin, Okyay et al. 2007) and in India(Gokhale, Kanade et al. 2004) withholding colostrum was associated with increased prevalence of stunting. In India (Kumar, Goel et al. 2006), delaying initiation of breast-feeding to more than six hours after birth, deprivation of colostrum and improper complementary feeding were significant risk factors for underweight child but wasting was not significantly associated with any one of these infant feeding practices. In a small case-control study conducted in rural village in India time of initiation of breast feeding and withholding of colostrum were not significantly associated with child underweight (Sanghvi, Thankappan et al. 2001). In a cross-sectional survey conducted in a food surplus region of Ethiopia, the main contributing factors for stunting in under-five year old children were the child being of a male gender, an increase in child's age, diarrhoea episode, withholding of colostrum, shorter duration of breastfeeding, pre-lacteal feeds, type of food, age of introduction of complementary feeding and method of feeding (Teshome, Kogi-Makau et al. 2009). In Uganda (Engebretsen, Tylleskar et al. 2008) sub-optimal infant feeding was a predictor of infant growth. However, investigators in Burkina Faso (Thiombiano-Coulibaly, Rocquelin et al. 2004) and in India (Sanghvi, Thankappan et al. 2001) did not replicate the negative effect of sub-optimal feeding upon child growth.

We are unaware of any previous study to have assessed the effect of antenatal CMD on early infant feeding practices. However, there is some evidence from LAMIC settings (Nhiwatiwa, Patel et al. 1998; Inandi, Elci et al. 2002; Rahman, Iqbal et al. 2004; Collin, Chisenga et al. 2006) that antenatal CMD frequently persists into the postnatal period. When measured late in pregnancy, antenatal CMD may be considered a proxy measure for early postnatal maternal mental health. In an economically disadvantaged environment, postnatal CMD is associated with deficits in a face-to-face mother-infant interactions (Murray and Cooper 1997), a compromised mother-infant relationship (Cooper, Tomlinson et al. 1999) and adverse child developmental millstones (Murray and Cooper 1997). In developing country settings, therefore, continuity of antenatal CMD into postnatal period may combine with, exhaustion as the result of labour to undermine the mother's ability to respond to her newborn's needs right from the start. Hence, it is reasonable to hypothesize a significant negative association between antenatal CMD and early infant feeding practices.

1.6.3 The effect of perinatal CMD on infant undernutrition

Maternal CMD is highly prevalent in LAMICs (Prince, Patel et al. 2007) and recent studies indicate a potential aetiological role in infant undernutrition (de Miranda, Turecki et al. 1996; Patel, DeSouza et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Rahman, Lovel et al. 2004; Harpham, Huttly et al. 2005; Adewuya, Ola et al. 2008; Stewart, Umar et al. 2008; Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009; Surkan, Kennedy et al. 2011). Infancy is a critical time for the well-being of the newborn which depends largely on the quality and quantity of care received from the primary caregiver (Rahman, Harrington et al. 2002), usually the mother. Postnatal CMD can affect the mother's mental and physical availability to the infant and thus compromise parenting quality (Cooper, Tomlinson et al. 1999; Rahman, Harrington et al. 2002). A meta-analysis of 19 studies conducted in high-income countries found postnatal depression to have a moderate-to-large adverse effect on maternal-infant interaction during infancy (Beck 1995). These findings have been replicated in South Africa, with depressed mothers exhibiting less sensitive engagement with their infants (Cooper, Tomlinson et al. 1999) resulting in an increased insecure attachment in the infants (Tomlinson, Cooper et al. 2005). Maternal depression was also associated with a less

stimulating home environment for infants in Jamaica (Baker-Henningham, Powell et al. 2003)

Maternal CMD might lead to infant undernutrition through a variety of mechanisms (Rahman, Harrington et al. 2002; Patel, Rahman et al. 2004). When present during pregnancy, maternal CMD has been associated with an elevated risk of low birth weight (Copper, Goldenberg et al. 1996; Patel and Prince 2006; Ferri, Mitsuhiro et al. 2007; Rahman, Bunn et al. 2007), which in turn is associated with infant undernutrition(Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004). Postnatal CMD may lead to early cessation of breastfeeding (Patel, DeSouza et al. 2003) or compromised hygienic feeding practices putting the infant at risk of infectious illnesses (Rahman, Bunn et al. 2007).

Studies from South Asia (Patel, DeSouza et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Rahman, Lovel et al. 2004; Black, Baqui et al. 2009) have consistently found postnatal CMD to be associated with infant undernutrition after adjusting for potential confounders. However, in Latin America findings have been more mixed, with maternal CMD associated with child under-nutrition in a cross-sectional community sample from Brazil (de Miranda, Turecki et al. 1996; Surkan, Kawachi et al. 2008), but not in a clinic-based study from Jamaica (Baker-Henningham, Powell et al. 2003) or a large population-based sample in Peru (Harpham, Huttly et al. 2005). A similarly inconsistent picture is emerging from sub-Saharan Africa (Harpham, Huttly et al. 2005; Collin, Chisenga et al. 2006; Tomlinson, Cooper et al. 2006; Adewuya, Ola et al. 2008; Stewart, Umar et al. 2008). In clinic-based studies from Nigeria (Adewuya, Ola et al. 2008) and Malawi (Stewart, Umar et al. 2008), maternal postnatal CMD was associated with infant undernutrition; However, in a population-based cross-sectional sample of two to 18 month old children in Ethiopia (Harpham, Huttly et al. 2005) and a population-based cohort in South Africa (Tomlinson, Cooper et al. 2006), no significant associations were noted between maternal CMD and child undernutrition. In another clinic-based study in Zambia (Collin, Chisenga et al. 2006) maternal postnatal CMD at six weeks was not associated with weight-for-age and length-for-age z scores of infants at the age of four months. In high income countries postinatal CMD does not seem to have any major effect on infant growth (Grote, Vik et al. 2010)

Methodological issues may explain some of the conflicting findings across studies in LAMICs. Variation in the age of children at nutritional assessment, heterogeneity of study participants across studies, rural versus urban setting, cultural validity of instruments used to ascertain maternal CMD, use of different nutritional indices as outcomes, as well as different scales of measurement (binary or continuous), the frequencies of exposure and outcomes investigated, the timing at which the effect of exposure on the outcome was evaluated, and the quality of study design may all play a part (Stewart 2007; Black, Baqui et al. 2009). Furthermore, the majority of published studies fail to take into account the potential impact of maternal CMD in pregnancy upon infant under-nutrition, mediated through low birth weight. Studies from LAMICs have tended to show that the prevalence of maternal CMD is higher in pregnancy than in the postnatal period, underlining the importance of examining the impact of antenatal CMD. Only one study, from Pakistan, has evaluated the effect of maternal CMD in pregnancy on child nutritional status prospectively using a population based cohort (Rahman, Iqbal et al. 2004) and showed that CMD in pregnancy significantly compromised the nutritional status of infants at six and twelve months of age. In sub-Saharan Africa, health service coverage is generally low (Lawn, Cousens et al. 2005; UNICEF 2008; WHO 2010) which means that clinic-based studies are examining a selected population; this may lead to bias, since women who seek help because their child is under-nourished and ailing may be more likely to be psychologically distressed.

1.6.4 Modelling the effect of perinatal CMD on infant growth

1.6.4.1 Current practice of modelling the effect of perinatal CMD on infant undernutrition

Modelling the effect of maternal mental ill-health on physical growth of children commonly involves dichotomizing the main exposures determined from measures of psychological morbidity generated from symptom counts (Baker-Henningham, Powell et al. 2003; Patel, DeSouza et al. 2003; Rahman, Lovel et al. 2004; Harpham, Huttly et al. 2005; Stewart, Umar et al. 2008; Black, Baqui et al. 2009) or diagnostic instruments (Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006; Surkan, Ryan et al. 2007; Adewuya, Ola et al. 2008; Surkan, Kawachi et al. 2008) and use of various formulations of physical growth measurements as outcome measures. These growth outcome measures include standardised weight (Baker-Henningham,

Powell et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Rahman, Lovel et al. 2004; Harpham, Huttly et al. 2005; Tomlinson, Cooper et al. 2006; Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009), percentile of weight (Patel, DeSouza et al. 2003; Adewuya, Ola et al. 2008), standardised length (Rahman, Iqbal et al. 2004; Harpham, Huttly et al. 2005; Tomlinson, Cooper et al. 2006; Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009), and percentile of length (Patel, DeSouza et al. 2003; Adewuya, Ola et al. 2008). Scale of an outcome measures (e.g. continuous versus binary) and the validity of distributional assumptions made about the residuals are critical in selecting appropriate modelling strategies.

After dichotomizing standardized measurements of physical growth to generate the proportion of undernourished children several researchers have evaluated the effect of maternal CMD on these outcome measures at one (Patel, DeSouza et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Lovel et al. 2004; Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009) or at more than one time points (Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006; Adewuya, Ola et al. 2008) using logistic regression. Under this modelling strategy the expected value of the target growth measurement is assumed to be the sum of the linear combination of the predictor variables and the residuals on a logit scale rather than on its original scale:

$$\log it(p_i) = \ln\left(p_i/(1-p_i)\right) = \sum_{j=1}^{k} \beta_j x_j + \varepsilon_i$$

where p_i stands for the probability with which the binary growth outcome assumes one (i.e. the individual child is undernourished), k stands for the number of predictor variables (i.e. potential risk factors) that are include in the model, the residuals, ε_i , i =1,2,..n are discrepancies between the observed and model predicted logit for each individual. These residuals are assumed to be independently generated from the logistic function.

Investigators who have evaluated the effect of maternal mental ill-health on growth measurement at more than one time-point have used independent logistic regression models at each time-point where the outcome is defined. Rahman (Rahman, Iqbal et al. 2004) evaluated the effect of CMD on growth of children at six and 12 months of age. Tomlinson (Tomlinson, Cooper et al. 2006) investigated the effect of concurrent CMD

on growth of children at two and 18 months, and the effect CMD measured at two months on growth of children at 18 months. Adewuya (Adewuya, Ola et al. 2008) used logistic regression at six weeks, three months, six months and nine months to evaluate the effect of postnatal depression on child growth in the first year of life. In all of these studies the effect of maternal CMD on growth trajectory was interpreted from the consistency of the findings at various time points obtained from cross-sectional analyses.

Some investigators have combined continuous with binary growth outcome (Tomlinson, Cooper et al. 2006) and others have only modelled continuous growth outcome measures (Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Black, Baqui et al. 2009) using linear regression with the assumption of normally distributed residuals to investigate the effect of CMD on child psychical growth. With this modelling technique the expected value of the outcome variable y (i.e. $E(y) = \mu_{y/x}$) is assumed to be equal to the sum of the linear combination of k predictor variables and random residuals:

$$E(y_i) = \mu_{y_i/x} = \sum_{j=1}^{k} \beta_j x_j + \varepsilon_i$$

where k stands for the number of predictor variables (i.e. potential risk factors) that are include in the model, the residuals ε_i are deviations of the model predicted mean value from the observed growth measure. The residuals are assumed to be normally and independently (NID) distributed, with a mean of zero and constant variance ($\varepsilon_i \sim NID(0, \sigma^2)$).

One of the key assumptions required for the validity of linear regression analysis is independence of observations. This assumption is violated under cluster sampling or repeated measurement of individual subjects over time. If responses are generated from individuals clustered in some meaningful ways, for example, geographical area, household or villages, or within individual over time so that an outcome measure is highly correlated within the cluster compared to individuals outside the cluster, the assumptions required for linear regression are no longer valid. The same is also true when we are dealing with binary outcome measures.

Recognizing the problems of using logistic regression and linear regression when the assumption of independence of observations is violated researchers in the field of

perinatal CMD and child growth in LAMICs have started to use modelling techniques that take account of within-cluster correlation of growth measurements (Harpham, Huttly et al. 2005; Surkan, Ryan et al. 2007). In a situation where the interest was to adjust for the effect of selected covariates without measuring their individual effect, Surkan and colleagues (Surkan, Ryan et al. 2007) used a multilevel model while investigating the effect of maternal psychological morbidity on child growth. Investigators in a multi-country study (Harpham, Huttly et al. 2005) have also used generalized estimating equations (GEE) (Twisk 2003) to model the effect of CMD on child growth while adjusting for the clustering nature of individual growth measurements within geographical settings. Both methods aim to take account of inter-individual dependency while estimating the effect of predetermined risk factors on given outcome variables. The GEE method defines the working covariance structure for within individual/cluster dependence of observations and uses this covariance structure in the estimation process to adjust the standard error of model parameters simultaneously. However, the multilevel modelling technique defines separate parameters within the domain of the model parameters to account for the inter-individual/cluster dependence and estimates these parameters simultaneously with other model parameters (Twisk 2003).

1.6.4.2 Current practice of modelling the effect of perinatal CMD on infant growth trajectories

A few recent studies in high income countries have used longitudinal modelling techniques (Ertel, Koenen et al. 2010; Ertel, Koenen et al. 2010; Grote, Vik et al. 2010) to evaluate the effects of maternal CMD on child growth trajectories. In a European cohort (Grote, Vik et al. 2010), postnatal CMD did not significantly affect trajectories of length, weight-for-length, and body mass index of children in the first 24 months of life. However, postnatal CMD was inversely and significantly associated with the trajectory of weight, which remained constant in the first year of life and decreased thereafter (Grote, Vik et al. 2010). In the USA antenatal and postnatal CMD seems to act differently in the first three years of child life. Maternal postnatal CMD was significantly associated with obesity at three years of age (Ertel, Koenen et al. 2010) and with an increased height-for-age and longer leg length at three years of age (Ertel, Koenen et al. 2010). However, antenatal CMD was significantly associated with a deficit in weight-for-height (Ertel, Koenen et al. 2010) and with an increased leg length but did not

significantly affect height-for-age (Ertel, Koenen et al. 2010). A recent cross-sectional study among low-income households in Brazil found significant association between postnatal CMD and being overweight (i.e. above 95th or 85th percentile of weight-for-height) of children aged 6-24 months (Surkan, Kawachi et al. 2008).

Available evidence about the effect of CMD on growth trajectories in LAMICs is based on the findings from modelling cross-sectional associations of CMD and growth at several consecutive ages of children (Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006; Adewuya, Ola et al. 2008). In Pakistan (Rahman, Iqbal et al. 2004) the effect of antenatal CMD and persistent CMD on child growth was investigated at six and 12 months of age and complemented with graphical description of mean growth trajectory in the first year of life. The findings of this study were that (a) the adjusted effect of antenatal CMD on underweight and stunting were significant at both time points but the effect size on both outcomes were larger at six months than at 12 months, (b) the negative effects of persistent CMD on both growth outcomes were stronger than the negative effects of antenatal CMD at both time points and (c) CMD had a stronger negative effect on underweight than on stunting. In Pakistan and elsewhere in the region poverty and compromised maternal empowerment have been proposed as the likely common causes of higher prevalence of CMD and undernutrition, and the reported strong associations between the two factors (Patel, Araya et al. 1999; Patel, DeSouza et al. 2003; Rahman, Iqbal et al. 2004; Harpham, Huttly et al. 2005; Rahman, Bunn et al. 2007; Rahman and Creed 2007).

In South Africa (Tomlinson, Cooper et al. 2006), the effect of postnatal CMD (a) measured at two months on concurrent growth measurements and on 18 month growth measurement and (b) measured at eighteen months after delivery on concurrent growth measurements, were investigated using three different logistic regression models (for dichotomous outcomes) and three different linear regression (for continuous outcome) models. This study could not establish a significant effect of postnatal CMD on infant growth in the first 18 months of life. Although CMD measured at two months postnatal was marginally associated with a reduced weight-for-age z-score before taking account of any confounding that association was not significant after adjusting for birthweight. In Nigeria (Adewuya, Ola et al. 2008) the effect of postnatal CMD measured at six weeks of child birth was assessed on growth of infants at six weeks, three months, six

months and nine months using logistic regression. Although the associations reported in that study were not adjusted for any potential confounding factors, postnatal CMD had a significant effect on underweight and stunting of infants at three and six months. The effect size for both growth outcomes was stronger at six months compared to at three months of age, and was also stronger for underweight compared to stunting at six months of age.

1.6.4.3 Hypothesized pathways in the current literature on how perinatal CMD might affect child growth

Pregnancy, child delivery, child nurturing and child feeding exercises are gifts from nature given to females (the first two exclusively) and they are the only means that human generation guaranties its continuity. The process of creating a new life and to be fully responsible for its wellbeing is a demanding task and requires commitment, skill, energy, and dedicated time from the mother. People get involved in this natural activity with full consent and hoped to generate happiness out of the whole process. It is also evident that this natural process creates stress of various degree to the whole family, especially, to the mother who is the primary actor within the process. Hence, it is interesting to know whether maternal stress during perinatal period has significant negative effect on the well being of the child and if so what possible mechanisms might be involved in this interaction.

Inspired by the findings of previous studies showing that CMD is more common among women than men (Patel, Araya et al. 1999) and particularly common among women of childbearing age (Kumar 1994), the negative effect of postnatal CMD on child development (Murray and Cooper 1997) and by the preliminary indication of the negative effect of postnatal CMD on infant underweight in Brazil (de Miranda, Turecki et al. 1996) it was hypothesized that there could be a strong effect of perinatal CMD on infant growth in settings where the environment is not conducive for successful child rearing in terms of economic resources, hygiene and healthcare availability (Rahman, Harrington et al. 2002). The effect of antenatal CMD on the fetus mediated in various ways, direct impact of persistent or incident postnatal CMD on parenting, and frequency of negative life events associated with maternal CMD were hypothesized as possible

mechanisms by which infant growth could be compromised (Rahman, Harrington et al. 2002).

The hypothesized negative effect of perinatal CMD on child growth (Rahman, Harrington et al. 2002) was evaluated in different LAMIC using various study designs, involving different study populations, using various types of data analysis techniques, with different levels of adjustment for confounding effects and reported various conclusions (Patel, Rahman et al. 2004; Stewart 2007). There are several studies which showed that antenatal CMD is a risk factor for low birthweights in LAMIC (Patel and Prince 2006; Rahman, Bunn et al. 2007) although some studies did not replicate this finding (Collin, Chisenga et al. 2006; Hanlon, Medhin et al. 2009). In high income countries there is strong evidence supporting the negative effect of prenatal CMD on pregnancy outcomes including low birthweight (Grote, Bridge et al. 2010). However, the knowledge of the exact mechanism is not yet fully understood except for some studies showing that maternal biochemistry during pregnancy might be playing mediating role of the association (Field, Diego et al. 2004; Field, Diego et al. 2006).

A recent review (Stewart 2007) which has synthesized the available literature on the effect of perinatal CMD on child growth and development with emphasis on the former summarized the mechanisms of how the negative effect of perinatal CMD might be manifested on child physical growth Figure 1.1

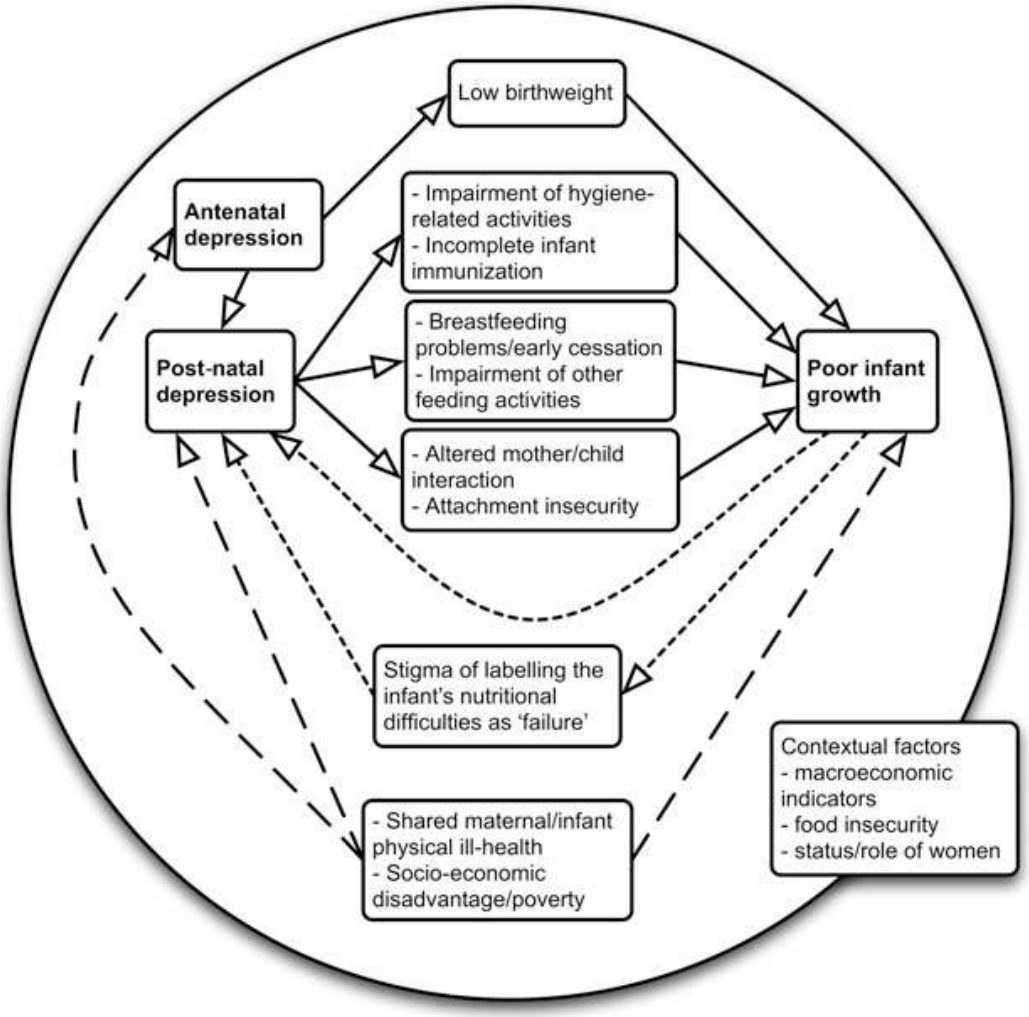

Figure 1.1: Mechanisms hypothesised in the literature of LMIC on how perinatal CMD might affect child growth and summarized by Stewart (Stewart 2007)

Subsequently published studies from LAMIC provided supportive evidences for some of the hypothesized mechanisms (Rahman, Bunn et al. 2007; Rahman and Creed 2007; Stewart, Umar et al. 2008; Black, Baqui et al. 2009), replicated some of the previous findings (Santos, Santos et al. 2010) or provided new highlight on the timing CMD to which growth is sensitive (Santos, Matijasevich et al. 2010). In agreement with previous cohort study in the UK which reported weight deficit up to four months of age (Wright, Parkinson et al. 2006) a recent cohort study from Europe reported that perinatal CMD only affects the weight of infants and that this effect declines after one year of age (Grote, Vik et al. 2010). The findings of a recent meta-analysis showed that CMD in

LAMIC is a significant risk factor of child underweight and stunting (Surkan, Kennedy et al. 2011)

1.6.4.4 Current practice of handling missing data, mediation, effects modifiers, clustering and confounding in modelling the effect of perinatal CMD on infant growth

Handling of missing data

Although it was not explicitly stated in the analysis section of the published papers the majority of researchers who have investigated the effect of CMD on attained growth or growth trajectories in LAMICs, with studies referenced in the previous pages have treated missing data using a list-wise deletion approach. However, in one study missing data were substituted with the mean of the non-missing observations before modelling the association (Stewart, Umar et al. 2008). Mean substitution and list-wise deletion of missing data assume that the data is missing completely at random (MCAR). Both methods produce a biased parameter estimates if the MCAR assumption is violated and list-wise deletion significantly affects power of the study if the percentage of missing data is large (Enders 2001). Mean substitution underestimates the variance and increases the probability of type I error (Schafer and Graham 2002)

In studies that have used multilevel modelling in LAMICs (Surkan, Ryan et al. 2007; Surkan, Kawachi et al. 2008) and in high income counties (Ertel, Koenen et al. 2010; Ertel, Koenen et al. 2010) to evaluate the effect of CMD on growth trajectories model parameters were estimated using full information maximum likelihood (FIML). This method uses all available information in the data to estimate parameters rather than reducing the sample size by dropping any case with missing data points or replacing the missing observations by some estimates. An alternative to FIML would be to use the multiple imputation (MI) technique which has been shown to perform equally well (Newman 2003; Abraham and Russell 2004). The assumption about missing data required by FIML and MI is missing at random (MAR) which is less restrictive than the assumption of MCAR.

Mediators

Although researchers investigating the effect of perinatal CMD on infant growth in LAMICs have discussed several possible mechanisms about how CMD might influence the foetus, pregnancy outcomes and child growth (Rahman, Harrington et al. 2002; Patel, Rahman et al. 2004; Stewart 2007) hypothesised mediating factors have not been investigated quantitatively (de Miranda, Turecki et al. 1996; Patel, DeSouza et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Rahman, Lovel et al. 2004; Patel and Prince 2006; Tomlinson, Cooper et al. 2006; Rahman, Bunn et al. 2007; Surkan, Ryan et al. 2007; Adewuya, Ola et al. 2008; Stewart, Umar et al. 2008; Surkan, Kawachi et al. 2008; Santos, Santos et al. 2010). In a few instances mediating effects of some variables in LAMICs were evaluated by the degree of attenuation of the coefficient for CMD and growth after including the hypothesized mediating variable into the model (Harpham, Huttly et al. 2005; Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009). Others (Santos, Santos et al. 2010) have suggested potential mediating factors in their concluding discussion without attempting to evaluate their roles in a multivariable analyses. Researchers in the USA added potential mediator variables into a fully-adjusted model and assessed the magnitude of attenuation observed in the magnitude of the coefficient of CMD (Ertel, Koenen et al. 2010). Other investigators have treated mediator differently as recommended previously (MacKinnon, Lockwood et al. 2002)

Effect modifiers

Some investigators (de Miranda, Turecki et al. 1996; Anoop, Saravanan et al. 2004; Harpham, Huttly et al. 2005; Surkan, Kawachi et al. 2008; Santos, Santos et al. 2010) addressed the issue of effect modification while assessing the effect of CMD on child growth in LAMICs by including the interaction of CMD and hypothesized effect modifiers into the model and evaluating the statistical significance of the resulting coefficient. In India (Anoop, Saravanan et al. 2004) birth weight and maternal intelligence were hypothesized as effect modifiers of the association between maternal CMD (current and postnatal) and child underweight. The finding of this study was that birthweight was a significant effect modifier of the effect of current CMD on underweight and maternal intelligence was a significant effect modifier of the effect of postnatal CMD on underweight. In a multi-country study (Harpham, Huttly et al. 2005),

child sex and age did not modify the effect of CMD on stunting and underweight. In Brazil, maternal age and number of children in the family were significant effect modifiers of the association between CMD and child underweight in one case-control study (de Miranda, Turecki et al. 1996). However, in another case-control study in Brazil (Santos, Santos et al. 2010) there was no significant effect modifier of the association between CMD and wasting. Similarly, maternal self efficacy of parenting was not a significant effect modifier of the association between CMD and stunting evaluated in a cohort study in Brazil (Surkan, Kawachi et al. 2008).

Confounding

With few exceptions (Tomlinson, Cooper et al. 2006; Adewuya, Ola et al. 2008), researchers investigating the effect of CMD on child growth in LAMICs have considered several potential confounding variables using logistic regression or linear regression depending on the scale of growth outcome measure under investigation. The analysis plan of a multi-country study (Harpham, Huttly et al. 2005) clearly identified which variables were hypothesized as potential confounding variables and which variables were hypothesized as mediating variables. Confounding variables frequently considered by researchers conducting these studies in LAMICs include infant gender, infant age, infant feeding practices (eg. duration of exclusive breast feeding, timing of the initiation of supplementary feeding), maternal characteristics (eg. age, nutritional status, education, empowerment, health) and some indicators of socioeconomic status.

Clustering of growth outcomes

The effect of clustering was considered in a few studies. A multi-country study (Harpham, Huttly et al. 2005) considered the clustering nature of growth of children within geographical settings and the investigators used GEE method to model the association while adjusting for the clustering. To adjust for the clustering of growth measurements due to neighbourhoods and interviewer styles in two studies in Brazil investigators (Surkan, Ryan et al. 2007; Surkan, Kawachi et al. 2008) used mixed effects modelling technique and considered the two variables as random effects. In a recent meta-analysis different studies were considered as random sample from all possible

studies and mixed effects model was used to estimate the pooled effect of CMD on child nutritional effect in LAMICs (Surkan, Kennedy et al. 2011)

1.6.4.5 Limitations of modelling techniques used to assess the effect of CMD on infant growth in LAMIC and the way forward

In terms of data analysis techniques the evidence we have so far reviewed regarding the relationship between maternal mental disorder and infant growth in LAMICs has been mainly based on standard modelling techniques varying from cross-tabulation to multiple logistic regression and linear regression. In recent years there have been some attempts to clearly specify possible mechanisms as much as the data allows them (Harpham, Huttly et al. 2005) and to use more advanced statistical modelling techniques that takes account of clustering of outcome measures in modelling the effect of CMD in growth in these countries (Harpham, Huttly et al. 2005; Surkan, Ryan et al. 2007; Surkan, Kawachi et al. 2008; Surkan, Kawachi et al. 2008). The majority of the researchers have discussed the possible mechanism in reference to previous evidence or have hypothesized plausible mechanisms to be addressed in future research. Although the focus of interest on the relationship of perinatal CMD on infant growth is more towards LAMICs, some research in the field in developed countries have recently started using modern advanced statistical modelling techniques as well (Ertel, Koenen et al. 2010; Ertel, Koenen et al. 2010; Grote, Vik et al. 2010).

Advanced modelling techniques like Path Analysis (PA) or more general Structural Equation Modelling (SEM) are more appropriate under some conditions (Bollen 1989; Vasconcelos, Almeida et al. 1998; Kline 2005). For example, when reciprocal causal effects are meaningful- for instance maternal CMD affects child nutrition and child nutrition may in turn influence maternal CMD, or some variables are allowed to serve as risk factors as well as outcome variables in system of equations that should be solved simultaneously, or latent variables are involved in modelling processes by allowing two or more measured indicators for the theoretical construct. Investigators have explored the potential advantages of using these statistical methods over standard modelling techniques (Sheehan 1998; Vasconcelos, Almeida et al. 1998; Feldman, Dunkel-Schetter et al. 2000; Field, Diego et al. 2004; Ross, Sellers et al. 2004). PA allows investigation of more complex models, providing information that could have been previously

overlooked, such as how the interrelations among independent variables in a model affect the dependent ones (Vasconcelos, Almeida et al. 1998). Unlike standard regression modelling techniques they can distinguish between potential confounders and meditating factors. Findings like the effect of mothers' antenatal biochemistry (Field, Diego et al. 2004) and social support (Feldman, Dunkel-Schetter et al. 2000) on birth weight, the indirect effect of biological variables on antenatal mood change (Ross, Sellers et al. 2004) and identification of the pathways through which different stressors affect birthweight (Sheehan 1998) were obtained from the application of SEM. However, PA and SEM are not the only advanced modelling techniques and the selection of a modelling technique should be based on the specific research question and availability of appropriate data.

When the interest is to evaluate change over time (eg. growth of children) and predictors of that change, multilevel growth modelling (Bryk and Raudenbush 1987) and latent growth modelling techniques (Meredith and Tisak 1990; Curran and Hussong 2003) are obvious alternatives to traditional logistic regression and linear regression. Both modelling techniques take account of inter-individual correlation while evaluating the effect of potential risk factors on growth (eg. perinatal CMD). The two modelling techniques were developed from two different perspectives (multilevel growth modelling from the perspective of mixed effects and latent growth modelling from a SEM perspective) and they use time of growth measurement in the model differently.

1.7.1 Structural Equation Modelling (SEM)

Structural Equation Modelling (SEM), also known as covariance structure analysis, covariance structure modelling, or analysis of covariance structure is a collection of related statistical techniques designed to model complex relationships between characteristics under investigation (Kline 2005). SEM gives an opportunity to use comprehensive methods for the quantification and testing of theories of complex relationships, to explicitly take into account the measurement error, and to use latent variables as a cause and as an outcome. The central hypothesis in SEM is that the covariance matrix of the observed variables can be expressed as a function of a set of model parameters (Bollen 1989) , i.e.

$$\sum = \sum(\theta)$$

where

$\sum$ is the population covariance matrix of the observed variables,

θ is a vector of model parameters, and

$\sum(\theta)$ is the covariance matrix written as a function of θ.

Modelling techniques such as correlation analysis, regression analysis, path modelling, and factor analysis can be obtained from full SEM by introducing restrictive assumptions. The complexity of these models is inversely related to the number of assumptions required by each modelling techniques (Bollen 1989; Kline 2005). The hierarchical links between some of these modelling techniques are well described (Musil, Jones et al. 1998) and extensively discussed elsewhere (Kline 2005).

1.7.1.1 Description of selected concepts

The fact that SEM is a collection of interrelated techniques developed over a very long period of time in various disciplines (Bollen 1989) has resulted in several terminologies, some of them overlapping and others referring to distinct concepts. For meaningful discussion it is helpful at this stage to describe the meaning of concepts that will be referred to subsequently. These include

- Discrepancy indicators (residuals, error terms, unique factor, measurement error, random error)
- Class of variables (measured, latent, exogenous, endogenous,)
- Effects (direct, indirect and total),
- Recursive and non-recursive models,
- Association measures (i.e. correlation and covariance),
- Regression
- Path diagram.
- Path coefficient and path regression coefficients.

Residual, unique factor, error term, measurement error, and random error:
These terminologies are used to represent discrepancies between the true value and its estimate in a given investigation. In most instances investigators use residuals, unique

factors and error terms interchangeably to mean the same thing in different problem areas. Each term can be categorized as measurement error (also called systematic error) or random error.

Residuals or error terms are mostly attached to a model and they represent the portion of an outcome variable which is not accounted for by the specified model. For example, in a simple linear regression model, $y_i = \alpha + \beta * x_i + e_i$, the residual term (i.e. e_i) represents the influence of any other variable on the outcome variable yi but not due to x_i, because the influence of xi is already accounted for. In other words any variability in y which is not explained by x is taken care of by the residual term. This will mean that the effects of all unmeasured variables which might have an effect on y, measurement errors that might have occurred in recording values of y and random variations in y are all aggregated in one error tem. In this situation it is not possible to disaggregate these three sources of variability.

Unique factor is the concept developed in connection with factor analysis and it has the same meaning in factor analysis as that of residual in regression analysis (Browne, MacCallum et al. 2002). For example, if we are interested to measure socioeconomic status, many candidate indicator variables can be proposed but none of them is a perfect measure. In this particular example monthly income, education and current savings might be considered as indicator variables. Some of the variability in each of these three indicators is explained by the underlying unobserved common factor (i.e. socioeconomic status) and the remaining portion of the variability in each of the indicator variable is explained by some other factors. The latter part of the model (i.e. the portion which is not explained by the indicator variables) is called the unique factor and, as can be seen, plays the same role as the residual in the regression model.

Measurement error stands for the discrepancy between the true value and its estimated value due to the unreliability of measurements used to generate the estimate. On the other hand *random errors* are deviations from actual values that occur just purely by random process or by chance and cannot be predicted. For example, while recording current savings in a survey some people will report household savings, others will report the savings of the household head and others might report a reduced or increased amount

compared to their actual savings. In using this variable as one indicator of socioeconomic status we have three sources that can explain the variation in current savings: (1) the underlying shared factor for educational level, current saving and monthly income (i.e. socioeconomic status), (2) differential reporting behaviour of individuals (i.e. measurement error) and (3) some other factors other than socioeconomic status and measurement error (i.e. random error). In an ideal condition, the population value can be measured without any error as long as we are able to control measurement errors. However, measurements in research settings are done beyond ideal situations and we always expect an error just because of random fluctuations. This part of error is known as random variation or random error and it occurs in a random ways which can not be predicted within the process of measurement under investigation.

Latent and observed variable

In SEM the meaning of the concept "latent variable" or "factor" is crucial and needs clear understanding. In real life people use many concepts which are difficult to measure directly and only their proxy characteristics can be recorded. These theoretical or hypothetical constructs that have no direct operational method for measuring them or a precise method for assessing their degree of presence are known as latent variables or factors (Bollen 1989; Kline 2005). These theoretical concepts are created by researchers to fully understand the research area under investigation (Bentler 1980). The scores measured directly, in contrast to the latent variables that have to be measured through a proxy measurement, are termed observed variables. Observed variables serve as indicators of the underlying construct or latent variable. For example, manifestations of a construct (or latent variable) like antenatal CMD, domestic violence during pregnancy, and socioeconomic status can be observed by recording or measuring specific features of each concept in a particular environment. The recording or measurement of the construct is typically obtained by using pertinent instruments (e.g. tests, scales, self-reports, inventories, or questionnaires). Once constructs are assessed using carefully selected indicator variables with acceptable psychometric properties (Bollen 1989; Hoyle and Smith 1994) SEM can be used to test the plausibility of hypothetical assertions about potential interrelationships among the constructs as well as their relationships to the indicators or measures assessing them (Bollen 1989; Hoyle and Smith 1994).

Exogenous versus endogenous variables

One of the assumptions in SEM is that some variables have influence on other variables and these influences can be quantified in the process of model fitting. Those variables which receive influence from other variables within the context of the defined model are known as endogenous variables, whereas those variables which are not influenced by any variable in the specified model are known as exogenous variables. This is illustrated in **figure 1.2** that shows a hypothetical link between four variables.

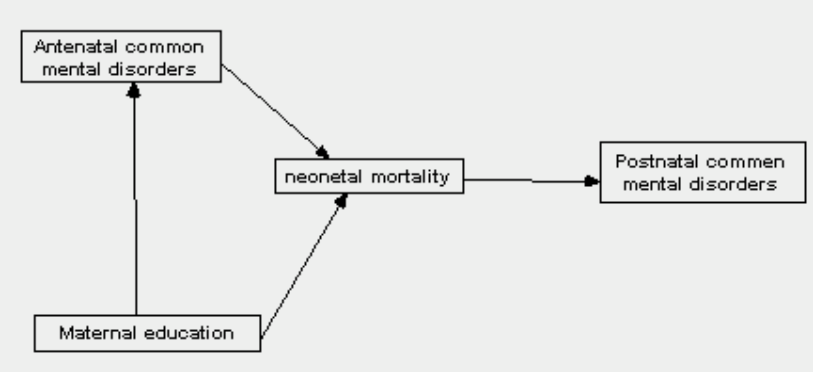

Figure. 1.2: Hypothesized link between maternal education, antenatal CMD, neonatal mortality and postnatal CMD

In figure 1.2 the only exogenous variable is maternal education, which is not influenced by any of the variables specified in the model. The other three variables, i.e. antenatal CMD, neonatal mortality and postnatal CMD are endogenous variables. In formulating regression type equations from the path diagram, like the one presented in figure 1.2, some endogenous variables can serve as independent variables in one equation and as dependent variable in another equation (eg. antenatal CMD and neonatal mortality in figure 1.2). However, exogenous variables can only serve as independent variables in one or more than one equation (eg. maternal education in figure 1.2).

Direct, Indirect and Total effects

Effect is a numerical measure that indicates the extent to which the value of a given outcome variable is changed due to a unit change in the value of the independent variable in a given network of relationships. Direct, indirect and total effects are defined with reference to one specific dependent variable and one specific independent variable in a

given network of relationships. Direct effect measures the unmediated influence of an independent variable on a dependent variable while indirect effect measures the influence of an independent variable through all possible mediating variables. Total effect is the sum of the direct effect and all indirect effects.

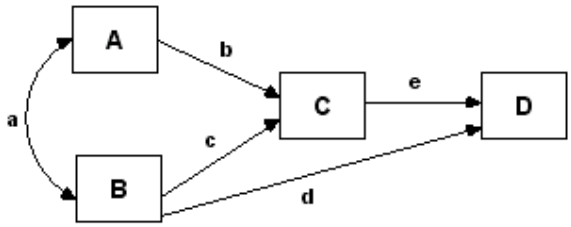

Figure 1.3 Hypothesized causal relationship between two exogenous (A, B), one mediating (C) and one outcome variables (D)

If we take the hypothetical relationship of four variables presented in figure 1.3 as an example where the capital letters represent measured variables and small letters represent path coefficients

- A and B are correlated in a non-causal manner (also called an unanalysed association to indicate that the reasons for the observed association are not investigated or ares not important to consider within the context of the model)
- A and B have a direct effect on C. Similarly, B and C have a direct effect on D
- C has a mediating role in the relationship of A and D as well as in that of B and D
- There is no direct effect of A on D
- A and B have an indirect effect on D individually through C (these effects cannot be measured in ordinary regression analysis).

If we take two variables, say A and D, from the example presented in figure 1.3 there is no direct effect of A on D. However, there are three indirect effects and the sum of these three indirect effects will constitute the total effect of A on D:

- One indirect effect through C, (its magnitude = b*e)
- Another indirect effect through B, (its magnitude is = a*d), and

- Another indirect effect through B and then though C (its magnitude = a*c*e)

Recursive and non-recursive models

The distinction between recursive and non-recursive model is based on the way variables are hypothesized to influence each other. In a recursive model one variable cannot influence a variable and at the same time be influenced by that variable in a given causal line. In a non-recursive model, however, variables are allowed to influence other variables (i.e. to be an independent variable) and at the same time to be influenced by the same variable (to be a dependent variable) in the same systems of relational equations (i.e. reversed causality). The hypothetical path diagram presented in figure 1.4 is an example of a non-recursive model as there are two reversal causalities (i.e. causal lines a and b as well as causal lines c and d). If we remove one causal line from each group (eg lines a and d) without affecting any of the other existing links of the network then the model would be changed into a recursive one. Distinguishing between recursive and non-recursive models has implications for the way the model is fitted to the data (Bollen 1989).

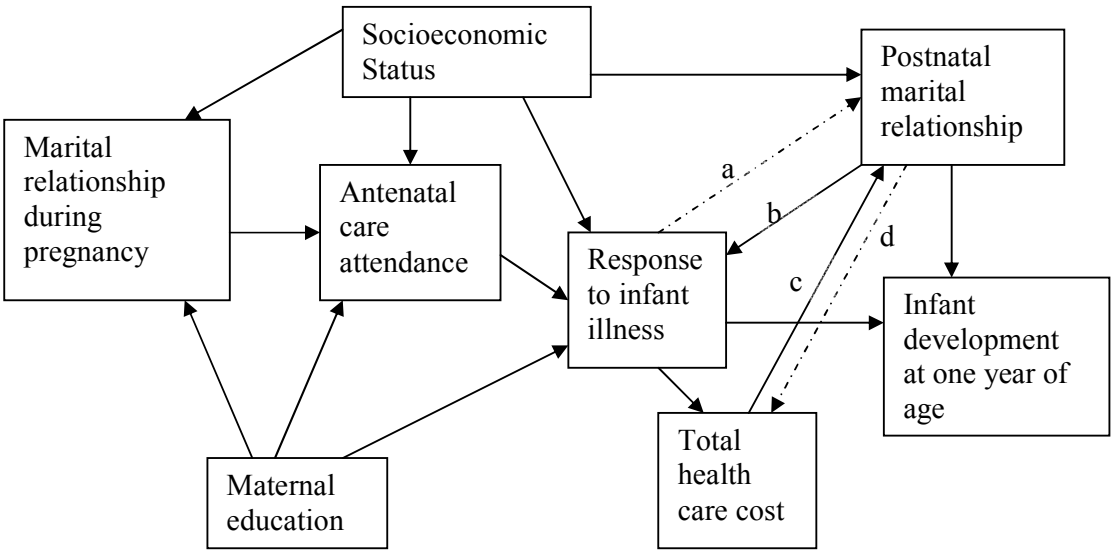

Figure 1.4: Hypothesized non-recursive associations between eight mother-child related characteristics

Correlation and Covariance

These are measures of non-directional relationships between two measured variables and they play a central role in SEM. When the two variables are continuous and normally distributed, the numerical value of the Pearson correlation coefficient (r) is obtained after standardizing the covariance of the two variables under investigation (Bollen 1989). For example, if we record information from n number of subjects for two numerical characteristics, say X and Y, then the Pearson correlation coefficient (r) is calculated as follows:

$$r = \frac{Cov(X,Y)}{\sqrt{Var(X)Var(Y)}} = \frac{\sum_{i=1}^{n}(x_i-\bar{x})(y_i-\bar{y})/(n-1)}{\sqrt{\sum_{i=1}^{n}\frac{(x_i-\bar{x})^2}{n-1} * \sum_{i=1}^{n}\frac{(y_i-\bar{y})^2}{n-1}}}$$

Hypothesising a cause and effect relationship is not required to model association of two variables using correlation analysis since this measure is symmetrical. If the two variables are highly correlated knowing the value of one variable will tell us the likely value of another variable. Depending on the nature of variables analysed different types of correlation coefficients are used in SEM. The following are examples of Pearson type and non-Pearson type correlation coefficients (Kline 2005)

A) Pearson type – involves either categorical or ordinal variables

- Point-biserial correlation – measures an association between binary variable and continuous variable
- Phi coefficient – measures an association between two binary variables
- Spearman's rank order correlation (Spearman's rho) – measures an association between two ordinal variables

B) Non-Pearson type that assumes underlying continuous and normally distributed latent variable for categorical variable(s)

- Biserial correlation – measures an association between continuous variable and binary variable
- Polyserial correlation – generalization of biserial correlation when the categorical variable has more than two categories
- Tetrachoric correlation – measures an association between two binary variables assuming that there is an underlying continuous and normally distributed latent variable for each of the two binary variables
- Polychoric correlation - generalization of tetrachoric correlation when the two variables have more than two categories each

Regression analysis

This is a modelling technique that requires specification of the cause and effect relationship among measured variables whose association is to be modelled and is hence more advanced than correlation analysis. If we have one outcome variable, say y, which is normally distributed, another variable, say x, measured without any error, normally distributed and linearly related to the mean value of the outcome variable, then the regression line that links y with x is given as $E(y/x_i) = \mu_{y/x_i} = \alpha + \beta * x_i$. Regression coefficients α and β, which are the parameters of the specified model, are estimated from sample data using ordinary least squares method. Regression modelling is the building block for structural equation modelling (Bollen 1989; Kline 2005). The independent variable(s) is(are) assumed to be measured with no error but the outcome variables have random errors which take all the variation not accounted for by the independent variables (Huba and Bentler 1982; Musil, Jones et al. 1998). Regression

coefficients that quantify the relationship between the outcome variable and each of the independent variables are interpreted as measures of change in the mean value of the dependent variable as the result of a unit change in the independent variable. When standardized values of measured variables are used in regression modelling the intercept takes the value of zero and the unit of measure for the regression coefficient becomes the standard deviation (i.e. we get standardized regression coefficients instead of usual regression coefficients).

Path diagrams

These are graphical methods used to present theoretically hypothesized causal and non-causal relationships among characteristics of interest (Wright 1934). In the application of SEM path diagrams are the preferred ways of presenting hypothesized interrelationship of variables. In this diagrammatic approach of presenting systems of equations, universally accepted specific symbols are attached to specific characteristics of SEM and they are the easiest ways to communicate the complex structural relations (Bollen 1989; Byrne 2001). Squares or rectangles are used for observed variables and circles or ellipses are used for latent variables (or unobserved true values). For example, in figure 1.5 all the residuals denoted by e_i, socioeconomic status, antenatal care attendance, child development at 12 months and undernutrition at six months are all latent variables and the remaining variables presented within rectangular blocks are observed variables.

Hypothesised functional relationships are represented graphically in a path diagram by one-way arrows (eg: association between socioeconomic status and child development at 12 months of age) and two-way arrows (eg. relationship between the unexplained variation in self report and institutional record of antenatal attendance of study subjects) (Wright 1923). One-way arrows are usually represented by straight lines, with arrow-heads at the end of the straight lines. The variable at the end of the arrow is assumed to be the effect and the one at the beginning is assumed to be the cause. Two-way arrows (sometimes referred to as two-ways paths and graphically represented as curved lines with an arrowhead at each end) are used to represent covariation between two variables (un-analysed association) and indicate that there is an association between the connected variables that is not assumed to be directional.

The idea of establishing causation using path diagram and path coefficients was debated from the very start of the concept of path diagrams as the method of analysis (Niles 1922) and that is still the case. Three conditions should be evaluated to establish a causal relation between variables – isolation, association, and direction of causality (Bollen 1989). While association and direction of causality may be fairly easy to examine, it is quite difficult to ensure that a cause and effect have been isolated from all other influences. For this reason, many researchers consider SEM models and the causal relations within the model as approximations to reality that can never really be proved (Bollen 1989). They can only be disproved or disconfirmed.

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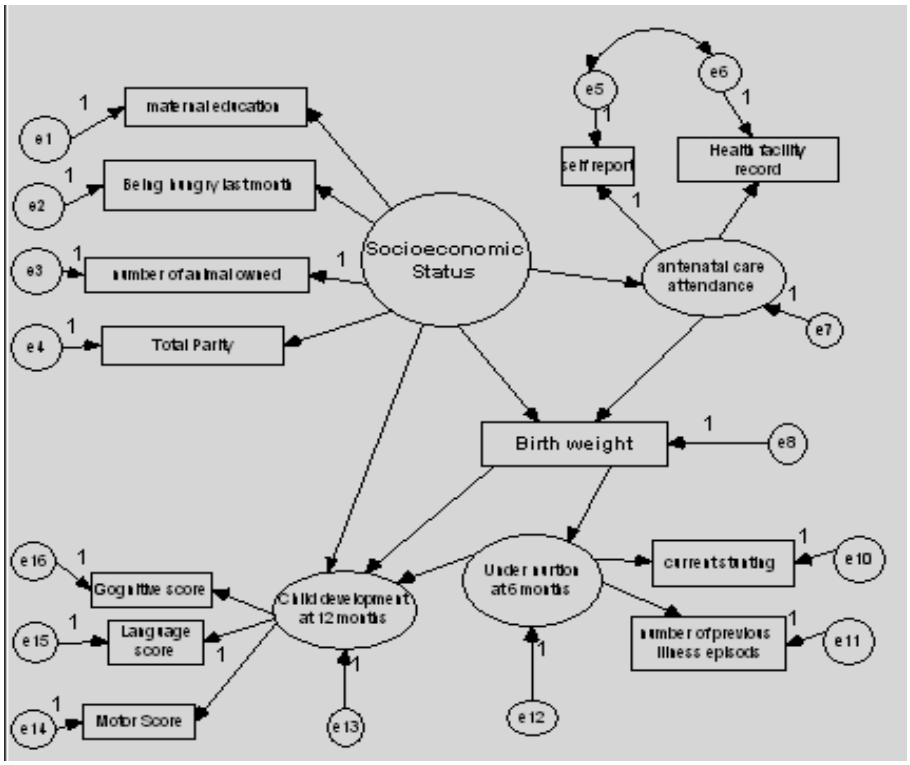

Figure 1.5: Path diagram of hypothesized association between selected mother-infant characteristics: some measured and others latent variables

1.7.1.2 Types of SEM

As previously described, the term SEM refers to a collection of modelling techniques that have been developed over a long period of time in different professions. Some of the techniques are more complicated than others. The more complicated approaches are generalizations of those simpler modelling techniques. In this subsection we will be looking into path modelling, factor analysis, full SEM, and structural regression modelling and latent growth modelling

A) Path analysis or path modelling

This modelling technique was introduced into the literature as early as 1918 by Sewall Wright in his generic work (Wright 1918) and fully described in the early 1920s (Wright 1920; Wright 1921; Wright 1923). The first critic on the validity of path analysis as a scientific method to establish a cause and effect relationship was published in 1922 (Niles 1922). The inventor of path analysis responded to the critique and advanced the methodology for decades (Wright 1923; Wright 1934; Wright 1960; Wright 1960), becoming the base for development of SEM (Wolfle 2003).

In its early stage of development, path analysis was meant to be a flexible means of decomposing the correlation of variables into systems of equations that described functional relations among them (Wright 1923; Wright 1934; Bollen 1989). For a long period of time its application was mainly in human genetics which might have benefited from the research interest of the inventor of path analysis and mathematical complexity of the technique as it was presented to other professions (Duncan 1966). Its application was then extended to different fields including education, and sociology starting in the second half of 1950s (Wolfle 2003).

With its increased application, several misunderstandings, both of purpose and of procedures, were noted and they became reasons for detailed assessment of previous applications of path analysis and clarification of what is possible and what is not possible in path analysis (Wright 1934). In this publication Wright re-iterated the basics of producing an appropriate path diagram, raised the importance of using standardized

symbols in path analysis, clarified ways of quantifying effects (i.e. direct, indirect and total), discussed interpretation of these effects and highlighted means of estimating the paths which may account for a set of observed correlations. Arguments made on the superiority of unstandardized path coefficient over standardized path coefficient (Turner and Stevens 1959) were explained by showing how the two measures of effect complement each other in the interpretation of the research problem (Wright 1960). Extending the application of PA in a recursive models Wright also investigated the application of PA to non-recursive models (Wright 1960). In his entire research work Wright underlined the fact that path coefficients are not intended to deduce causality just from the values of the correlation coefficients but to combine the quantitative information with such qualitative information as may be at hand on causal relations to give a quantitative interpretation (Wright 1934; Wright 1960)

PA extends the idea of regression modelling and gives the flexibility of quantifying indirect and total causal effects in addition to the direct effect which was also possible in regression analysis (Bollen 1989). In other words, regression analysis allows an independent variable to influence an outcome variable only directly but path analysis gives more flexibility and predictor variables are allowed to influence the outcome variable directly as well as indirectly through other mediating variables (Wright 1934). PA shares the following principles of regression analysis: (1) the direction of influence in the relationship of variables should be specified from the theory behind the investigation, (2) independent variables are assumed to be measured without error, (3) the relationship between target variables is linear and (4) any outcome variable in the system of equations under investigation has an error term attached to it.

There are three interrelated components in PA (Bollen 1989):

- translation of a conceptual problem into a pictorial presentation (i.e. path diagram described above) which shows the network of relationships,
- obtaining systems of equations that relate observed correlation and covariance to parameters (i.e construction of a system of simultaneous equations described implicitly and explicitly in the path diagram), and
- decomposition of effects of one variable on the other (i.e. direct, indirect and total) from the correlation of measured variables.

The pictorial presentation or path diagram helps clarify what is meant by the conceptually formulated problem and leads to formulation of systems of mathematical equations that can be solved to give estimates of effects knowing the correlation or covariance of measured variables. From the start the path diagram was advocated as the easiest tool to conceptualize the "causal relationship" as well as to decompose the correlation between different variables into different source (Wright 1920; Wright 1921; Wright 1923)

Figure 1.6 shows a path diagram of conceptually defined hypothetical associations between selected maternal and child characteristics (i.e. Maternal education (MED), Socio-economic status (SES), Domestic Violence during pregnancy (DV), Antenatal Common mental disorders (ACMD), Birth Weight (BW)). In this path diagram BW is an end point, MED is exogenous variable, and the remaining three are intermediate variables in the described association. In this network of associations a direct effect of MED is proposed on three outcome variables, namely, BW, ACMD, and SES. Other hypothesized direct effects are from DV to BW, from SES to BW, from SES to DV and ACMD, and from ACMD to BW. The indirect effect of maternal education on birth weight goes through four different routes, namely (1) through ACMD, (2) through SES then through ACMD, (3) through SES and then through DV, and (4) through SES status then DV then ACMD. The four residuals attached to the four outcome variables are independent from each other and also independent from each of the explanatory variables included in the network of association described in the path diagram.

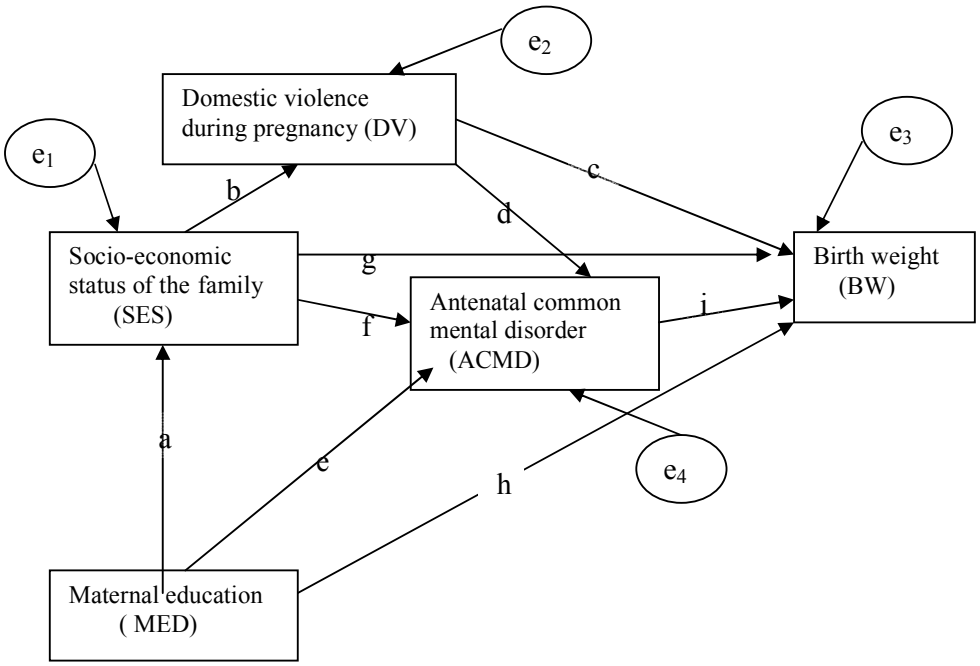

Figure 1.6: Path diagram describing hypothetically conceptualized association between selected characteristics of mothers and their new born

The associations presented in the path diagram in figure 1.6 can be translated into the following systems of equations:

$$BW = h*MED + i*ACMD + c*DV + g*SES + e_3$$

$$ACMD = e*MED + f*SES + d*DV + e_4$$

$$DV = b*SES + e_2$$

$$SES = a*MED + e_1$$

In setting these equations there are *a priori* assumptions that all measured variables to which the arrow head arrow points are influenced by the variables from which these arrows start. For example, in the last equation socioeconomic status of the household (SES) is affected by maternal education (MED) but the reverse is not the case. However, all variability that might exist in SES of the household will not be explained by maternal education and that part is handled by the residual term.

The following assumptions about residuals are also assumed in the path diagram by not showing their link via two-way curved arrows that indicates association:

COV (MED, e_3) = 0; COV(ACMD, e_3) = 0; COV(DV, e_3) = 0; COV(SES, e_3) = 0

COV (MED, e_4) = 0; COV(DV, e_4) = 0; COV(SES, e_4) = 0; COV(SES, e_2) = 0;

COV (MED, e_1) = 0; COV (e_1, e_2) = 0; COV(e_1, e_3) = 0; COV(e_1, e_4) = 0

COV (e_2, e_3) = 0; COV(e_2, e_4) = 0; COV(e_3, e_4) = 0

After translating path diagrams into systems of mathematical equations, the parameters of these equations (known as path coefficients) are estimated from the known values of the association measures between the measured variables (i.e. the correlation or covariance of the five measured variables) using one of the available estimation methods, for example, Ordinary Least Squares (OLS), Maximum Likelihood (ML), or Generalized Least Squares (GLS). After estimating path coefficients, the constructed path diagram can be traced by following appropriate tracing rules to get direct, indirect and total effects of the variable of interest both for recursive and non-recursive models (Wright 1934; Wright 1960).

In the 1950s and 1960s the application of PA was extended from just using observed variables in recursive models to include latent variables and non-recursive models, and moved from being an analytic tool of genetic research to be applied to a range of different disciplines (Goldberger 1972; Bentler 1980). Methodological advances made during this time (Simon 1954; Wold 1956; Wright 1960) became an incentive for its wider use in sociological research (Blalock 1961; Boudon 1965; Duncan 1966). This, together with the increasing accessibility of computers, helped PA to become part of a broader framework of sophisticated modelling techniques for examining causal associations, occurring towards the end of 1960s (Joreskog 1969; Wolfle 2003).

B) Factor Analysis

The objective behind factor analysis is to describe associations of variables in terms of a few underlying causes of these associations using the row data or variance co-variance matrix. This multivariate statistical technique has two main types, exploratory factor analysis (EFA) with a history of more than a century (Lovie and Lovie 1993), and confirmatory factor analysis (CFA) which was popularized starting in the mid 1960s (Brown 2006)

Exploratory factor analysis

EFA is a multivariate statistical technique which was invented by Spearman (Spearman 1904; Lovie and Lovie 1993) in 1904 and advanced within the field of psychology over the century (Spearman 1904; Lawley and Maxwell 1962). Spearman (Spearman 1904) used this technique for the first time in its current form with its current meaning and it is now widely used in the fields of social and behavioural sciences. The statistical methodology behind EFA was not properly investigated for more than 40 years which might have contributed to the controversies surrounding the application of EFA (Vincent 1953; Lawley and Maxwell 1962; Lovie and Lovie 1993; Fabrigar, Wegener et al. 1999). Regardless of how the model is fitted to the observed data, EFA attempts to identify causes of variability within each set of variables and to evaluate the contribution of each cause based on information available in the correlation between sets of measured variables (Vincent 1953). From the start, psychologists used EFA to reduce a large set of variables meant to measure related constructs into a smaller number of groups which are called constructs, factors or latent variables (Spearman 1904; Lawley and Maxwell 1962).

If we investigate *n* subjects and denote measurements made on these subjects by x_1, x_2, …, x_p then EFA assumes that there are a few common factors (less than *p*) underlying these measured variables, say *m*, and another group of unobserved (latent) variables each attached to each one of the observed variables and which uniquely affects that specific observed variable. In his original paper Spearman claimed that there is only one underlying factor for human ability and it is possible to measure this factor from different test results using EFA (Spearman 1904). Mathematically an EFA model with m underlying common factors can be expressed as follows:

$$x_i = \sum_{k=1}^{m} l_k f_k + e_i$$

where

l_k represents weight of the k^{th} common factor,

f_k represents the k^{th} underlying common factor, and

e_i represents a unique factor to the i^{th} measured variable.

Hence, in EFA the observed variable (i.e. x_i) is treated as an outcome or dependent variable and the causes for the observed variable are the unobserved common factors (f_k, k = 1, 2, ..m) and unobserved unique factors (i.e. e_i,).

A technique related to EFA is principal component analysis (PCA) (Hotelling 1933) which is frequently used as a factor extraction method (Fabrigar, Wegener et al. 1999; Conway and Huffcutt 2003; Watson and Thompson 2006). PCA summarizes the information available in p measured variables into p independent components, by taking linear combination of original variables, where the k^{th} component is described mathematically as follows:

$$c_k = \sum_{i=1}^{p} a_{ik} x_i$$

where

c_k (k = 1, 2, …p) is the k^{th} component

a_{ik} (i = 1, 2, …p) is the weight of observed x_i for component k

x_i is the i^{th} measured variable

The components are extracted in a hierarchical manner in such a way that (a) all the components are mutually independent and (b) each component explains the largest proportion of the variability from the total variability yet to be explained.

There are theoretical differences between PCA and EFA (Lawley and Maxwell 1962; Floyd and Widaman 1995). Unlike EFA in which measured variables are considered as outcome variables and the underlying common factors as predictor variables PCA assumes components as outcome variables which are influenced by the observed variables. There is no random error term in the definition of PCA implying that all the variability is explained by the extracted factors. However, there is a unique factor for every observed variable in the EFA model. This unique factor accounts for the variability within the measured variable occurring due to measurement error and random error. In EFA the extracted factors can be correlated but all the extracted components in PCA are mutually independent and ordered in terms of the proportion of variation they explain. The extracted factors in the EFA are unobserved or latent variables but components in the PCA are linear combinations of the observed variables. In PCA it is meaningful to compute scores for scores for each component, but that is not the case in EFA. Although PCA was designed purely for data reduction and its statistical properties differ from EFA

(Lawley and Maxwell 1962), it has been considered as a simple substitute for EFA by some researchers (Lawley and Maxwell 1962; Floyd and Widaman 1995)

In the 1900s the application of EFA was mainly intended to evaluate organization of mental ability (latent variable) using some test results (measured variables) in a way that emphasised the relationship of latent variables and measured variables (Lawley and Maxwell 1962; Bollen 1989; Lovie and Lovie 1993). The extracted factors are allowed to be correlated without imposing cause and effect relationship between them. In other words, factors are only used as causes of observed variables and the methodology of EFA does not allow one factor to be the cause for another factor. As with the development of PA, the advancement of EFA was full of controversies (Smith 1950) and there are still some controversies regarding how it is implemented (Fabrigar, Wegener et al. 1999; Conway and Huffcutt 2003; Watson and Thompson 2006).

A review carried out by Bentler (Bentler 1980) and others (Steiger 1979; Mulaik 1986; Bollen 1989) of the historical development of EFA, shows that the method was advanced for about 60 years by psychologists without knowledge of the development of path analysis in the field of biometrics (Wright 1920; Wright 1921; Wright 1923; Wright 1934). Most of the terminologies used within EFA are somewhat different from other multivariate statistical methods. This might reflect the significant role of psychologists in the development and application of EFA for several years (Kendall 1950). After reviewing methodological details of path analysis and factor analysis Goldberger (Goldberger 1972) showed the relevance of these methods to the problems of econometrics and advocated for the importance of a unifying analytical framework

In the1950s and 1960s, statisticians (Simon 1954; Wold 1956), biometricians (Tucker 1958; Wright 1960; Wright 1960) and mathematicians (Mulaik 1986) contributed to the methodological development of path analysis and EFA. Although Wold (Wold 1956) did not explicitly addressed the statistical issues in relation to EFA, he made a comparison of experimental and observational studies and highlighted the importance of substantive theory in any attempt to make causal inferences from observational studies. Simon (Simon 1954) separated spurious correlation from genuine correlation of two variables by introducing the third variable into the system and showed how causal interpretation can be reached from correlation. Tucker presented statistical methodologies that showed

how the EFA technique could be used to estimate parameters of hypothesized causal relationships (Tucker 1958). Turner and his colleagues (Turner and Stevens 1959) described methods to get estimates of path coefficients from recursive and non-recursive path models using Ordinary Least Squares (OLS) and maximum likelihood (ML) methods. They also described the issue of model identification and advantage of using the methodology of EFA in conjunction with that of path analysis. Wright used his PA methodology (Wright 1960) to estimate path coefficients in a non-recursive models.

Methodological advances made by statisticians (Simon 1954; Wold 1956) and biometricians (Tucker 1958; Wright 1960; Wright 1960) in the 1950s and 1960s had the significant effect of encouraging sociological methodologists (Blalock 1961; Blalock 1963; Boudon 1965; Duncan 1966) to look into ways of combining the simplicity of path analysis with the idea of latent variables. Extending the work of Simon (Simon 1954), in which the third variable was used to differentiate between spurious and true causal correlation between two variables, Blalock (Blalock 1961) investigated causal association among five variables. Both investigators (Simon 1954; Blalock 1961) underlined the need for prior knowledge of the subject in order to arrange the variables in their causal order. Boudon (Boudon 1965) combined the method of separating spurious correlation from true causal correlation with the principle of path analysis (Wright 1934). Blalock extended his own previous work (Blalock 1961) on cause and effect among measured variables to include latent variables (Blalock 1963). Duncan (Duncan 1966) did not particularly focus on latent variables but introduced the application of path analysis and path coefficients (Wright 1934) to the sociological literature using re-analysis of previously published data.

In additional to the methodological advances of path analysis and EFA, accessibility of computer programs was a break through in sophistication of data analysis (Mulaik 1986). Application of EFA has significantly increased in different disciplines starting from the late 1960s (Bentler 1980). Following the introduction of computer programs like Linear Structural Relations (LISREL) and advances made in statistical theory in early 1970s, the idea of factor analysis was extended and hypothesis testing about factor structure became possible (Mulaik 1986; Bollen 1989; Floyd and Widaman 1995). The practice of EFA was upgraded to confirmatory factor analysis (CFA), enabling explicit hypotheses about factor structure, derived from theory or previous research, to be tested.

Despite its long history and wide application, EFA has always been accompanied by criticism regarding its contribution to theory development as well as the manner in which it is sometimes applied (Fabrigar, Wegener et al. 1999). Factor indeterminacy has always been the point of discussion (Steiger 1979; Mulaik 1986; Steiger 1994). After analysing artificial data with a known structure, Armstrong (Armstrong 1967) demonstrated his reservations as to the utility of factor analysis in the development of theory. Ehrenberg (Ehrenberg 1962) discussed the pitfalls of factor analysis by raising 10 different questions and classified factor analysis as a technique which is confused with no interpretative guidance of the final end-results. Use of EFA without proper methodological development (Kendall 1950; Lawley and Maxwell 1962) was one of the most common criticisms of EFA during the first half of the twentieth century (Ehrenberg 1962; Fabrigar, Wegener et al. 1999).

Supporters of EFA claims that the criticism of the manner in which EFA is sometimes applied arise from the decisions that researchers make when conducting the analysis (Fabrigar, Wegener et al. 1999). In EFA, researchers do make a number of important decisions with respect to how the analysis is performed and wise decision at this level can minimizes criticisms as to the validity of EFA results (Finch and West 1997; Brown 2006). There are at least five major methodological issues that a researcher should consider at this stage (Fabrigar, Wegener et al. 1999): (1) the number of variables to be included in the study and issues related to the sample (i.e. size and nature of the sample), (2) whether EFA is the correct analysis technique to answer the research objective at hand, (3) selection of a specific procedure to fit the model, (4) the number of factors to be retained after factor analyzing the data, (5) selection of a rotation technique to simplify interpretation of the factors. These points are detailed here after.

(1) number of variables to be included, sample size and nature of the sample: The stability of extracted factors in EFA is affected not only by sample size but also by the number of indicator variables per factor, the reliability of these indicator variables and the magnitude of their communalities (MacCallum, Widaman et al. 1999). Hence, an adequate number of indicator variables should be sampled from the domain of interest to come up with important factors, and variables that are irrelevant to the domain should not be included to avoid the risk of obscuring true common factors and/or emerging of

spurious common factors (Fabrigar, Wegener et al. 1999). A more accurate result might be obtained if multiple indicator variables with acceptable psychometric properties are used for one common factor (MacCallum, Widaman et al. 1999). There is some recommendation for the number of observed variables to be at least three to five times the expected number of common factors (Fabrigar, Wegener et al. 1999) although rules of thumb of this nature were not recommended after conducting simulation studies (MacCallum, Widaman et al. 1999). Several recommendations are available to determine sample size although they are full of drawbacks (Floyd and Widaman 1995; Fabrigar, Wegener et al. 1999; MacCallum, Widaman et al. 1999). If more indicator variables per common factor are included and on average the communalities are 0.7 or higher, sample sizes as small as 100 might be enough to give accurate estimates of population parameters. A sample size of at least 200 might be enough under moderate communalities, and the required sample size increases as the magnitude of communality goes down (Fabrigar, Wegener et al. 1999; MacCallum, Widaman et al. 1999)

(2) Is factor analysis appropriate choice for the given goal of the analysis?

Before using EFA, researchers should check that their research objective is to arrive at a more parsimonious conceptual understanding of a set of measured variables by determining the number and nature of common factors needed to account for the pattern of correlations among the measured variables. This is the key to prioritizing EFA over PCA, that latter being more appropriate if the objective of investigation is data reduction. The implementation of PCA involves taking scores on a large set of measured variables and reducing them to scores on a smaller set of composite variables that retains as much information from the original variables as possible without any attempt to model the structure of correlations among the original variables (Fabrigar, Wegener et al. 1999). However, it is a common practice to observe results from PCA being used to address the objective of EFA (Floyd and Widaman 1995; Fabrigar, Wegener et al. 1999; Conway and Huffcutt 2003; Watson and Thompson 2006).

(3) Selection of a specific procedure to fit the model

A review of the applications of EFA showed that researchers seem to depend too much on PCA to fit their models (Fabrigar, Wegener et al. 1999), although more appropriate application of EFA is increasing with time (Conway and Huffcutt 2003). Most commonly recommended estimation methods for EFA include principal axis factoring,

iterated principal factoring, and ML factoring. The latter provides a wide range of fit indexes for model evaluation which make it preferable (Fabrigar, Wegener et al. 1999; Conway and Huffcutt 2003). However, before using ML researchers should check if the assumption of multivariate normality is not severely violated otherwise ML can produce misleading results (Hu, Bentler et al. 1992; Curran, West et al. 1996). In case of non-normality of the data, remedial methods should be taken before extracting the factors using the ML method (Fabrigar, Wegener et al. 1999)

(4) The number of factors to be retained after factor analyzing the data

The goal of EFA is to determine the number of "major" factors underlying the measured variables and the decision of how many factors to retain is both statistical and substantive (Fabrigar, Wegener et al. 1999). Rules have been proposed to decide on the number of factors to retain, and their pons and cons debated (Zwick and Velicer 1986). Use of ML estimation has provided additional ways to decide on the number of factors to retain (Fabrigar, Wegener et al. 1999). The cost of retaining the incorrect number of factors can be substantial (Levonian and Comrey 1966; Wood, Tataryn et al. 1996). Hence, the decision on the number of factors to be retained needs a balance between parsimony (i.e. a model with few common factors) and plausibility (i.e. a model with a sufficient number of common factors to adequately account for the correlation among measured variables).

5) Selection of rotation technique to make interpretation of the factors simple.

If EFA results in a solution with more than one factor, then there are an infinite number of alternative orientations of factors in multidimensional space that will explain the data equally well. This implies the absence of a unique solution provided that we have a more than one factor solution. This problem has been debated for several years (Steiger 1979) and several criteria having been proposed to overcome the difficulty of interpreting the factor solutions (Weiss 1971). The most commonly used criterion to overcome the problem of indeterminacy is the property of simple structure (i.e. factors which are most easily interpretable, meaningful, and replicable) suggested in late 1940s and which can be obtained by rotating the factors in multidimensional space (Fabrigar, Wegener et al. 1999; Conway and Huffcutt 2003). Simple structure refers to the condition that each factor is defined by a subset of measured variables with high factor loadings, and each measured variable will have a high loading only on one factor and small loadings on others. Several rotation methods have been developed (Weiss 1971) and the utility of

these methods extensively researched (Fabrigar, Wegener et al. 1999). The main classification that can be made between these rotation methods is orthogonal (all the factors are uncorrelated) and oblique (factors are correlated). Varmax is the most popular orthogonal rotation method and it attempts to maximize the variance of squared loadings on a factor. Among oblique rotations, the direct quartimin rotation, promax rotation, and Harris-Kaiser orthooblique rotation are frequently used (Fabrigar, Wegener et al. 1999)

Confirmatory Factor Analysis (CFA)

In the late 1960s causal modelling techniques that were in use for decades in genetic research (Wright 1934), psychology research (Spearman 1904) and econometric research (Goldberger 1972) converged into a unified framework (Joreskog and Lawley 1968; Joreskog 1969) making it possible to explicitly hypothesize factor structure and test its fit with the observed covariance of the measured variables (Bentler 1980; Floyd and Widaman 1995). Beyond the chi-square test a significant number of fit indexes also became available to assess the extent of model fit (Hu and Bentler 1995; Hu and Bentler 1998; Hu and Bentler 1999).

Before the advancement of computer technology and introduction of confirmatory factor analysis (CFA) (it is also called restricted factor analysis) into the literature in the 1960s EFA was used extensively among psychologists as quantitative support in development and evaluation of theories and standard instruments (Comrey 1988; Floyd and Widaman 1995). In a research area where much knowledge is unavailable, EFA can be valuable as supportive guidance in making decisions (Mulaik 1987; Bollen 1989). However, the effort of consolidating theories or instruments (i.e. verification of hypotheses) on the basis of EFA was always challenged (Ehrenberg 1962; Armstrong 1967) and EFA does not have sufficiently flexible statistical properties to overcome the critics (Steiger 1979; Mulaik 1986; Hoyle 1991). For example EFA does not allow the following restrictions (Comrey 1988; Bollen 1989) which are now possible with CFA

- a priori hypothesis on the number of underlying common factors to retain,
- some common factors to influence a subgroup of the measured variables and other common factors to influence some or all measured variables,
- some common factors to be independent and others to be associated

- some residuals or unique factors to be correlated

Under the newly introduced unifying framework it became possible to extend the basic idea of EFA by relaxing some of the assumptions of the basic factor model which were required for the implementation of EFA, leading to development of CFA (Joreskog 1969). The focus of CFA, also known as measurement models (Bollen 1989), is exclusively with construct relations (O'Grady 1983; Tanaka and Huba 1984 ; Bollen 1989). It allows researchers to use their understanding of the theory and put forward an *a priori* hypothesis about the latent variable model and the measured variables that make up this model (Joreskog 1969). CFA enables the testing of hypotheses directly that could only be approached in indirect ways using EFA (Joreskog 1969; Cole 1987; Floyd and Widaman 1995).

An example of CFA with one underlying common factor and four indicator variables is helpful to describe some properties of CFA. A one common factor CFA model given in figure 1.7 illustrates that every measured variable is influenced by one common factor (i.e. marital satisfaction) and a second factor (ie. e_1, e_2, e_3, or e_4) that accounts for the variability unique to the measured variable. All variables presented in circles are latent variables and those presented in rectangular boxes are measured variables.

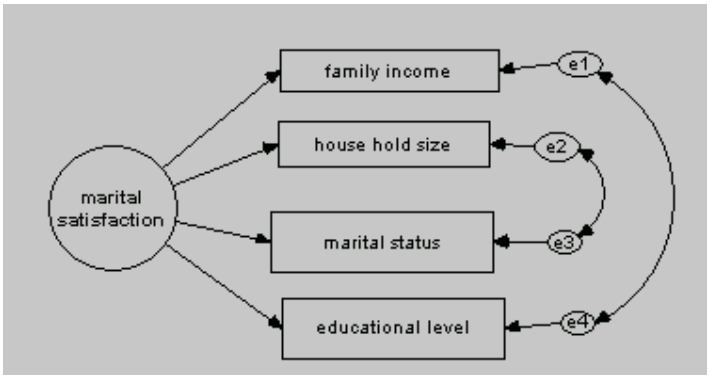

Figure 1.7: Path diagram describing the hypothesized link between marital satisfaction (i.e. latent variable) and four indicator variables

The following four measurement equations are evident in the path diagram:

$$x_i = p_i f + e_i \quad i = 1, 2, 3, 4$$

where

x_i (i= 1, 2, 3, 4) represents the i^{th} observed variable

f represents marital satisfaction factor and

p_i represents factor loading of the i^{th} measured variable

e_i represents unique factors

Some of the unique factors are correlated (i.e. e_1 with e_4 and e_2 with e_3) indicating the possible presence of common factor(s) they influence together outside the present model.

While using the basic common factor model, CFA relaxes some of the restrictive assumptions which were necessary for the implementation of EFA (Joreskog 1969). For example, in figure 1.7 some unique factors are allowed to be correlated but that would not be possible in EFA. In a given investigation if we measure p characteristics (say x_1, x_2, …, x_p) on n subjects and if we want to factor analyze the correlation matrix of these measurements with the hypothesis that there are k common factors (k <p) the basic factor model can be described in matrix form as follows (Joreskog 1969)

$$x = \Lambda f + e$$

where

- x is a column vector of the observed **p** measurements,
- *f* is a column vector of k (k < p) common factors,
- e is a column vector of p residuals, which represent the combined effect of specific factors and random errors, and
- $\Lambda = \left[\lambda_{ij} \right]$ is a p x k matrix of factor loadings.

The population covariance matrixes of x, *f* and e are denoted respectively by Σ, Φ and Θ. The expected value or mean of all the common factor *f* and all residuals e are assumed to be zero.

The CFA model described in figure 1.7 is the special case of a general CFA model with many underlying common factors. Hence, for the example CFA in figure 1.7 there is only one common factor and the parameters are as follows:

- The factor loading matrix is $\Lambda = \left(\lambda_1, \lambda_2, \lambda_3, \lambda_4 \right)'$,
- The variance of the common factor *f* is $\Phi = \phi$, and
- The variance-covariance matrix of four unique factors (i.e. e_1, e_2, e_3, and e_4) is

$$\Theta = \begin{pmatrix} \theta_1 & & & \\ 0 & \theta_2 & & \\ 0 & \theta_{32} & \theta_3 & \\ \theta_{41} & 0 & 0 & \theta_4 \end{pmatrix}.$$

- Sample variance and covariance matrix of the observed four variables (F for family income, H for household size, M for marital status and E for educational level) is the usual covariance matrix where the lower triangular matrix which could be presented as follows.

$$S = \begin{pmatrix} Var(F) & & & \\ Cov(H,F) & Var(H) & & \\ Cov(M,F) & Cov(M,H) & Var(M) & \\ Cov(E,F) & Cov(E,H) & Cov(E,M) & Var(E) \end{pmatrix}$$

Information available in the sample variance and covariance matrix of the measured variables will be used to estimate all unconstrained parameters of CFA model (i.e. factor loadings, factor variance, variance of unique factors and unconstrained covariance of unique factors).

CFA gives more control to researchers over the model building process and setting their hypotheses compared to what is possible in EFA (Joreskog 1969; Hoyle 1991; Hoyle and Smith 1994). For example some elements in the factor loading matrix Λ can be hypothesized to be zero based on theory or previous research unlike in EFA in which every element in the factor loading matrix is estimated from the data. CFA allows some of the unique factors as in figure 1.7 or all of the unique factors to be correlated. The number of common factors to retain from the analysis should be based on an *a priori* hypothesis in CFA but that is not the requirement in EFA. In CFA each latent variable included in the model is usually measured by its own set of observed indicators that are not influenced by other common factors described in the model and no specific directional relationships (i.e. dependent – independent type) are assumed between the latent variables, only that they are allowed to be correlated with one another. However, in EFA all indicator variables are allowed to load on every underlying common factor and simple structure is attained by rotating the extracted factors. In CFA it is possible to overcome the need for factor rotation and to attain simple structure by choosing the fixed parameters appropriately to have a desired property (Joreskog 1969).

The path diagram presented in figure 1.8 is an example of a CFA model with two correlated common factors (marital satisfaction measured twice - during pregnancy and after giving birth) and four indicator variables corrected for measurement errors (i.e. each indicator has a unique factor or residual attached to it). In this CFA model the following implicit assumptions are made:

- The two common factors are correlated and the magnitude of their correlation is not fixed but to be estimated from the data. This is shown by allowing the two common factors to be connected with a two-headed curved arrow and the arrow not being assigned any numerical value.
- There is no cross-loading of indicator variables on the two underlying common factors (i.e. one indicator variable can only load on or be influenced by one of the two common factors). This is guaranteed because a one-headed arrow is connecting to a given indicator variable from one underlying common factor but not from the other common factor

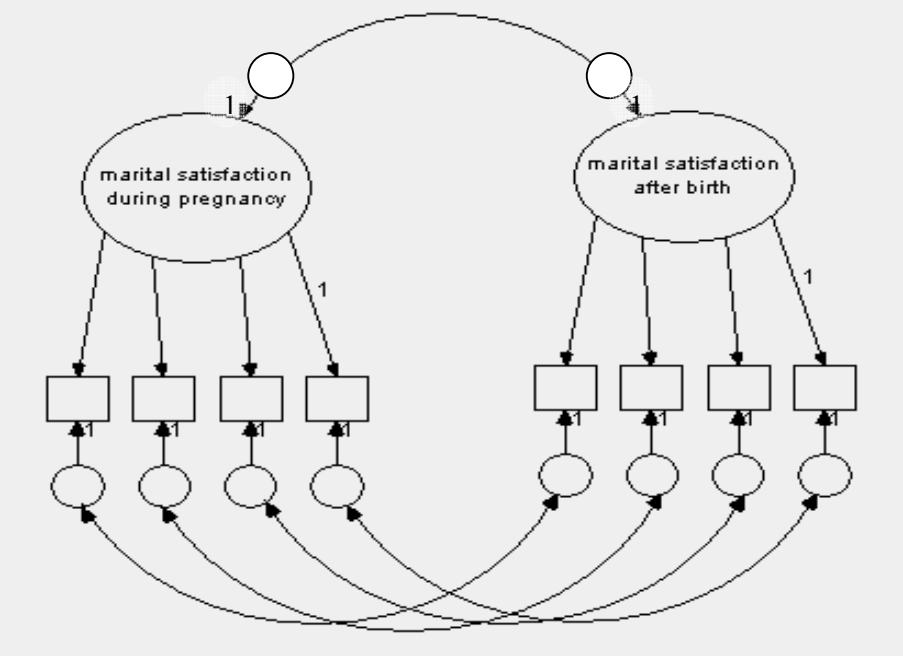

Figure 1.8: Example of two correlated CFAs with correlated unique factors

- The eight unique factors associated with eight measured variables are independent from the two underlying common factors but some unique factors are correlated. This is evident (a) by the absence of a two-headed arrow

connecting underlying factors and unique factors, and (b) the pair of unique factors being connected by a two-headed arrow.

- The scale for all latent variables (i.e. the two common factors and all unique factors) is determined by following unit loading indicator rules (Brown 2006). This is attained by assigning one to the loadings of unique factors and to the loadings of common factors with respect to one indicator variable each. Scaling the latent variables in this way results in unstandardized path coefficients. Another possible method of scaling the latent variables would be to fix the variance of each latent variable at one and allow all the loadings to be freely estimated from the data. This method of assigning a scale to latent variables results in standardized path coefficients.

With the above assumptions of the CFA model presented in figure 1.8, the factor loading matrix Λ, variance and covariance matrix of the unique factors Θ and the variance and covariance matrix of the two common factors Φ are summarized here in their respective order.

$$\Lambda = \begin{bmatrix} \lambda_{11} & 0 \\ \lambda_{21} & 0 \\ \lambda_{31} & 0 \\ \lambda_{41} & 0 \\ 0 & \lambda_{52} \\ 0 & \lambda_{62} \\ 0 & \lambda_{72} \\ 0 & \lambda_{82} \end{bmatrix}, \quad \Theta = \begin{bmatrix} \theta_1 & & & & & & & \\ 0 & \theta_2 & & & & & & \\ 0 & 0 & \theta_3 & & & & & \\ 0 & 0 & 0 & \theta_4 & & & & \\ \theta_{51} & 0 & 0 & 0 & \theta_5 & & & \\ 0 & \theta_{62} & 0 & 0 & 0 & \theta_6 & & \\ 0 & 0 & \theta_{73} & 0 & 0 & 0 & \theta_7 & \\ 0 & 0 & 0 & \theta_{84} & 0 & 0 & 0 & \theta_8 \end{bmatrix} \text{ and } \Phi = \begin{bmatrix} \phi_1 & \\ \phi_{12} & \phi_2 \end{bmatrix}$$

Because of the assumptions included in the CFA model as presented in figure 1.8, the values of the two factor loadings [i.e. $\lambda_{41}, \lambda_{82}$] and the loadings of all the unique factors are fixed to be unity. The absence of cross-loading of indicator variables on the two common factors is shown by the zero loadings within the factor loading matrix. Moreover, out of 36 non-redundant parameters of the variance and covariance matrix of the unique factors Θ, four covariance and eight variances are left to be freely estimated from the data and the remaining 24 covariance are fixed to be zero by the assumptions imposed on the CFA model.

Like any statistical method CFA should be considered as a complement rather than a substitute for good quality data (Martin 1982; Brown 2006). Before deciding to use CFA as the modelling technique and estimating model parameters, at least the following practical criteria must be checked (Cole 1987; Brown 2006)

- The sample size should be large enough to enable reliable estimation of model parameters. There are several recommendations on this and one rule of thumb is to consider 100 as the minimum, 200 to 300 as moderate and anything more than 400 to be considered as large. Another recommendation is to have at least 5 to 10 cases per parameter. However, the sensitivity of different parts of the model might be different to the sample size and model-based sample size estimation might be preferable to following rules of thumb
- If the plan is to use a correlation or covariance matrix as an input data for CFA then there should not be any missing data to avoid listwise deletion or case-wise deletion in generating the covariance matrix. However, if the decision is to use raw data to fit the CFA model missing data can be handled using full information maximum likelihood (FIML) or Multiple imputation (MI).
- Each latent variable should have multiple indicators, a minimum of two indicators per latent variable if there are more than two correlated common factors and a minimum of three indicators per latent variable if there is only one common factor. It is always better to have more indicators with acceptable reliability than having fewer indicators.
- Indicator variables for the underlying latent variable should be continuous and normally distributed. Severe violation of normal distribution assumption inflates overall goodness-of-fit measures and increases the risk of type I error (West, Finch et al. 1995). If categorical variables are among the indicator variables of the underlying latent variable, estimators other than FIML (eg. weighted least squares, robust weighted least squares) are more appropriate to fit the CFA model and other SEM (Muthen and Muthen 1998-2009)

Use of categorical variables as indicators of an underlying common factor involves the assumption of an underlying continuous, normally distributed latent variable for every categorical variable (Brown 2006). This implies the need for extracting appropriate variance-covariance matrixes (and mean vector if applicable) (Mehta, Neale et al. 2004)

before fitting the CFA model or any SEM. This point will be elaborated later in this chapter under the subheading "Latent growth modelling for binary outcomes".

C) Structural Regression Modelling

Structural regression modelling represents the convergence of independent research traditions in psychometrics(i.e. factor analysis or reliability theory), econometrics (i.e. simultaneous equations or error-in-variable models) and biometrics (i.e. path analysis) (Bentler 1980). From the late 1950s to late 1960s, researchers in the field of sociology (Blalock 1963; Duncan 1966) followed Simon's method (Simon 1954) of separating spurious correlation from true correlation in an attempt to establish causal relationships (Wright 1934), involving highly detailed analysis of data.

Investigation of underlying constructs using factor analysis (the "measurement model" in SEM terminology) and decomposition of the correlation of measured variables into different components using path analysis (the "structural model" in SEM terminology) were integrated into one general model as the result of statistical and computational advances attained in 1970s (Joreskog 1969) and implementation of this new modelling technique became easily accessible to applied researchers (Bentler 1980).

Mathematically, structural regression models (SRM) combine principles of CFA and PA (Bollen 1989). They are designed to evaluate an entire hypothesized multivariate model that includes hypothesized structural linkage (1) among latent variables and (2) between each latent variable and its respective indicator variables (Bollen 1989; Hoyle and Smith 1994; Musil, Jones et al. 1998). The logical structural regression model building process would be described as follows (Bollen 1989; Hoyle and Smith 1994; Kline 2005).

- to establish several measurement models that link observed variables to their corresponding latent variables equivalent to obtaining relevant confirmatory factor analysis models,
- to follow the principles of path analysis and relate the latent variables to formulate SRM equivalent to considering some latent variables as causes of other latent variables, and

- use a structural regression modelling technique to test the plausibility of hypothetical assertions about their explanatory relationships equivalent to estimation of model parameters using an appropriate fitting function

As described above, the structural regression model (SRM) has two main portions, namely, a measurement model and a structural model (Bollen 1989; Hoyle and Smith 1994). The relationships between the observed variables and the underlying latent constructs make up the measurement model portion of the equation. For example the path diagram presented in figure 1.9 shows one possible full SRM. In this diagram there are four measurement models defined by four latent variables (antenatal common mental disorder, socioeconomic status during pregnancy, antenatal common mental disorder at 12 month postnatal and socioeconomic status at 12 months postnatal) and their corresponding indicator variables. The structural part of the SRM is represented by arrows that link two or more latent variables in a causative way. For example in figure 1.9 the structure that remains after removing the indicator variables attached to the four latent variables (i.e. common mental disorder during pregnancy, socioeconomic status during pregnancy, antenatal common mental disorder at 12 month postnatal and socioeconomic status at 12 months postnatal) together with their respective unique factors, represents the structure model.

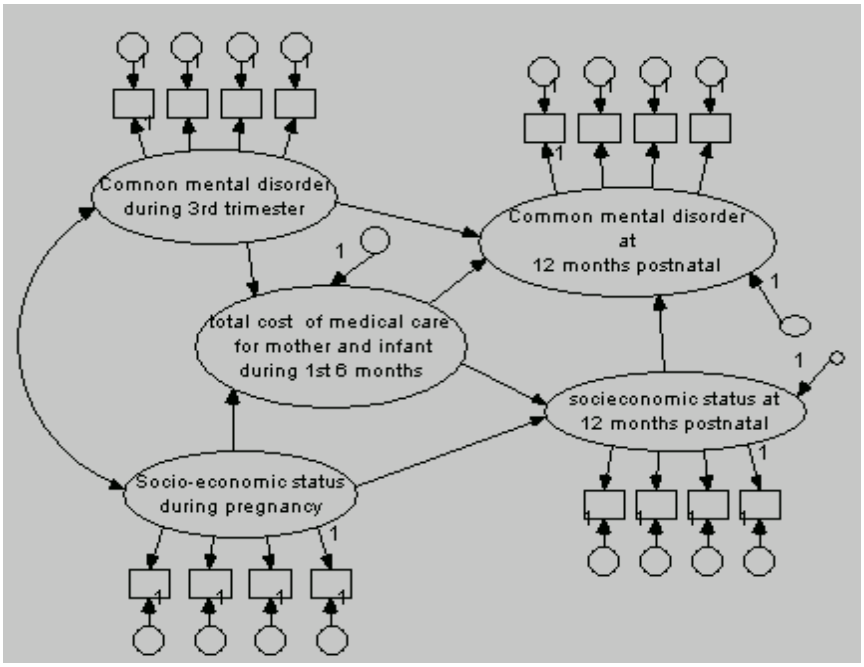

Figure 1.9: Path diagram giving an example of a structural regression model

In terms of their path diagrams, SRMs will always have at least one path (one-way arrow) leaving a putative explanatory latent variable and ending at another latent variable. For example in figure 1.9 there are

- three one-way arrows that end up on the latent variable "common mental disorder at 12 months postnatal",
- two one-way arrows that end up on the latent variable " socioeconomic status at 12 month postnatal", and
- two one-way arrows that end up on the latent variable "total cost of medical care for mother and infant during the first six months ".

Thus, in a way, structural regression models can be viewed as extensions of path analysis models except that, instead of being conceived in terms of only observed variables, the models also include latent variables (Hoyle and Smith 1994).

In SRM, a variable may be an independent variable in relation to a certain variable that follows it, a dependent variable, and a route for an indirect path (i.e. mediator) from other variables that precede it in the model. For example, in figure 1.9 "total cost of medical care for mother and infant in the first six months " is an independent variable in relation

to the two postnatal outcome variables but dependent in relation to the two antenatal latent variables. Similarly "socioeconomic status at 12 month postnatal" is a dependent variable in relation to the two latent variables but an independent variable in relation to another latent outcome variable.

To present SRM using matrix notation we should distinguish between the main components (structural part and measurement part) of the model and again between exogenous and endogenous variables. With this in mind and with the assumption that we have

- m latent endogenous variables (i.e. they are considered as outcome variables and their values will be determined within the process of model estimation),
- n latent exogenous random variables (i.e. their values are determined by conditions outside the model estimation process),
- n unique factors associated with each endogenous latent variable,
- p endogenous measured variables with associated measurement errors (i.e. they are considered as outcome variables within the model fitting process), and
- q exogenous measured variables with associated measurement errors (i.e.

it is possible to think of the following three matrix equations (Bollen 1989).

(a) Structural model:

$$\eta = B\eta + \Gamma\xi + \zeta \quad \text{and}$$

(b) Measurement model :

$$y = \Lambda_y \eta + \varepsilon \quad \text{for endogenous measured variables and}$$

$$x = \Lambda_x \xi + \delta \quad \text{for exogenous measured variable}$$

where

- $\eta =$ m x 1 vector of latent endogenous variables
- $\xi =$ n x 1 vector of latent exogenous variables
- $B =$ m x m coefficient matrix (path coefficients) showing the influence of latent endogenous variables on each other
- $\Gamma =$ m x n coefficient matrix (path coefficients) showing the influence of exogenous latent variables on endogenous latent variables

- ζ = m x 1 matrix of latent residuals, each element has an expected value of zero, and is independent from exogenous latent variables
- Λ_x = q x n matrix of coefficients and Λ_y = p x m matrix of coefficients, and measures the influence of the latent variable on its corresponding indicator variable in the appropriate matrix
- y = p x 1 vector of measured endogenous variables
- x = q x 1 vector of exogenous variables
- ε = p x 1 is vector of measurement errors with expected values of zero and assumed to be uncorrelated with ξ and ζ
- δ = q x 1 is vector of measurement errors with expected values of zero and assumed to be uncorrelated with ξ and ζ

By specifying the pattern of elements in each of the four coefficient matrices of the above equations (i.e. B, Γ, Λ_x and Λ_y), two covariance matrices of residuals ($E(\varepsilon * \varepsilon') = \Theta_\varepsilon$, $E(\delta * \delta') = \Theta_\delta$) and two covariance matrices of the latent random variables ($E(\eta * \eta') = \Phi$ and $E(\xi * \xi') = \Psi$), we can get simpler models including multivariate linear regression models, path models or structural equations with observed variables, or factor analysis models.

In the example presented in figure 1.9, there are three endogenous latent variables (m=3) (i.e. CMD at 12 months, socioeconomic status (SES) at 12 months, and medication cost (COST)), and two exogenous latent variables (n=2) (i.e. CMD during pregnancy and SES at pregnancy at baseline). This gives a dimension of 3-by-3 for the B matrix and a dimension of 3-by-2 for the Γ matrix of path coefficients. In the same example there are no exogenous measured variables implying that q = 0. However, there are 16 endogenous measured variables (p = 16), 16 unique factors associated with each measured endogenous variable and three latent error variables associated with three endogenous latent variables.

After specifying SRM on theoretical grounds the first step will be to establish all measurement models (Bollen 1989; Brown 2006) and the process then logically follows a model fitting process. Like previously described simpler models, assumptions made

about the variables (eg. normality of the indicator variables and underlying latent variables, scale of indicator variables, missing data mechanisms) and sample size will influence the choice of model fitting technique. In most practical conditions, originally hypothesized measurement models and SRM as a whole do not fit to the data under investigation. Under these circumstances modification of model should be based both on substantive theory and statistical significance of appropriate model parameters (Brown 2006)

D) Latent growth modelling

Latent growth modelling (LGM) is a statistical technique developed within the framework of SEM with the objective of modelling change over time and to identify predictors of that change (Meredith and Tisak 1990; Bollen and Curran 2006). The main focus of LGM is on patterns of growth, decline, or both, in longitudinal studies (e.g., on aspects such as initial status and rates of change over time) and enables researchers to examine both intra- and inter individual differences in patterns of change. It can also be used to examine the relationships between patterns of change and other background characteristics or time-dependent factors that might influence the change of interest.

Before the introduction of factor analysis in 1950s as the possible analysis tool for repeated measures (Baker 1954; Rao 1958; Tucker 1958) Analysis of Variance, Analysis of Covariance and Multivariate Analysis of Variance were standard procedures to model change over time (Wishart 1938; Bollen and Curran 2006). Following methodological advances, Baker (Baker 1954) factor analysed 20 repeated measures and extracted four underlying latent factors, and showed that different factors represent different stages of growth. Four years later Tucker (Tucker 1958) and Rao (Rao 1958) independently expanded the previous work on factor analysis and introduced the idea of using latent factors to estimate known functional forms relating time to the repeated measures. Anderson (Anderson 1963) discussed application of factor analysis from the perspective of multiple time series data. After several years it became possible to integrate the initiative of modelling repeated measures using EFA into the general SEM framework using the methodological advance of CFA over EFA (Meredith and Tisak 1990).

In using LGM to investigate change the first step is to establish a model that describes the magnitude and variability of individual growth trajectories (i.e. fitting an unconditional LGM). Like any other modelling technique, specification of subject level growth trajectories or unconditional LGM is influenced by the type of outcome at hand (e.g. continuous versus binary in this thesis). While specification of the functional relationship between time and the repeated measure is relatively easier for a continuous repeated measure, specification of the corresponding functional relationship of time with a binary outcome needs some modifications (Mehta, Neale et al. 2004; Bollen and Curran 2006). The magnitude and interpretation of growth parameters are also dependent on how time entered into the model (Biesanz, Deeb-Sossa et al. 2004) which will be briefly described below and in the methods chapter as applied to infant growth measures. After deciding on the best fitting subject level trajectory (i.e. unconditional LGM) the next step is to assess factors that might have significant effect on the parameters of unconditional LGM (i.e. fitting conditional LGM) (Meredith and Tisak 1990; Curran and Hussong 2003; Bollen and Curran 2006).

Unconditional LGM for continuous outcome

For a continuous growth outcome it has been shown that the unconditional growth trajectory model can be obtained from a CFA model after relaxing the following assumptions (Meredith and Tisak 1990):

- In a CFA model the underlying common factors have a mean of zero. However, in LGM the objective is to study how the mean trajectory of the target outcome changes with time and hence it is assumed to have a value different from zero.
- In CFA model indicator variables are assumed to have a mean of zero. This follows from the assumption that the mean of the underlying common factor as well as the mean of the unique factor is zero. However, in the LGM formulation the mean of the indicator variable is different from zero. In fact it is equal to the product of the mean of the underlying common factor and the corresponding factor loading.

To facilitate the use of ML to estimate model parameters and to be able to make inferences about growth trajectories (a) the assumption of normal distribution is imposed

onto the underlying common factors and unique factors which will then be implied for the indicator variables and (b) independence of underlying common factors and unique factors is assumed (Meredith and Tisak 1990). With these assumptions it is possible to write a CFA model that shows repeated measures as functionally related to the passage of time (Curran and Hussong 2003)

$$y_{it} = f(\lambda_t) + e_{it}$$

where

y_{it} is measure y for individual i at time t

λ_t is the value of time at t

$f(\lambda_t)$ reflects the functional relation between time and the outcome of interest

e_{it} is the residual for individual i at time t

The type of function $f(\lambda_t)$ we select to represent the functional relationship between time and the outcome of interest determines the shape of the growth trajectory. In LGM time is parameterized via the factor loadings that relate the repeated measures to the underlying latent factors (Meredith and Tisak 1990; Biesanz, Deeb-Sossa et al. 2004). Using the flexibility of CFA in deciding which factor loadings are fixed and which factor loadings are freely estimated (Joreskog 1969), it is possible to define different functional forms of growth.

For example if the pattern of a given data set shows a linear relationship with time the growth trajectory for individual i will have the following linear function which can easily be extended to a polynomial function of any degree (Curran and Hussong 2003):

$$y_{it} = \alpha_i + \beta_i \lambda_t + e_{it}$$

where

- y_{it} is measure y for individual i at time t
- λ_t is the factor loading on the linear slope of the underlying trajectory. For this linear growth function the values of λ_t will be fixed quantities directly linked to the time at which outcome y_{it} was measured.
- α_i is the intercept of the underlying trajectory for individual i
- β_i is the linear slope of the underlying trajectory for individual i
- e_{it} is the residual for individual i at time t

In the linear growth function described above, the intercept α and the mean β are allowed to vary from individual to individual and this is shown by including the subject specific index i. If we can assume that the study participants are random samples from the population then it is meaningful to consider individual subject specific parameters of the trajectory (α_i and β_i , i = 1, 2, …, n) as random variables and write them out as follows (Curran and Hussong 2003):

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i}$$

$$\beta_i = \mu_\beta + \zeta_{\beta_i}$$

Where

- μ_α is the mean intercept of the trajectory of all individuals
- μ_β is the mean slope of the trajectory of all individuals
- ζ_{α_i} is the deviation of each individual's intercept from the mean intercept $(\alpha_i - \mu_{\alpha_i})$
- ζ_{β_i} is the deviation of each individual's slope from the mean slope $(\beta_i - \mu_{\beta_i})$

After fitting unconditional linear LGM using ML to the observed repeated measures (e.g. infant length in this thesis) in any software suitable for SEM, it is possible to estimate mean intercept of the trajectory (i.e. μ_α) , mean slope of the trajectory (i.e. μ_β) , the variance of the intercept (i.e. σ_α^2), the variance of the slope (i.e. σ_β^2), the covariance of intercept and linear slope (i.e. $\sigma_{\alpha\beta}$), and time of measurement specific variance of the repeated measures (σ_t^2, t = 1, 2, …T assuming that there are T repeated measurement time points). If independence of residuals of the repeated measures is not assumed it is also possible to estimate hypothesized variance and covariance matricies of these residuals.

When it comes to the magnitude and interpretation of growth parameters (eg. μ_α and μ_β in the above unconditional LGM) the way the passage of time is coded has a significant role (Biesanz, Deeb-Sossa et al. 2004). If this phenomenon is not well understood it might affect how the model is structured and the growth parameters are

interpreted (Kurdek 1999; Biesanz, Deeb-Sossa et al. 2004; Bollen and Curran 2006). There are instances where the intercept could not have any real meaning in terms of the objective under investigation. For example, the mean length of infants when time is zero will not be meaningful. To overcome this potential problem of interpretation the common practice is to centre passage of time at some meaningful point (e.g. growth of infants at the age of two months) so that the intercept can be interpreted as the average value of the outcome of interest at that time-point (Biesanz, Deeb-Sossa et al. 2004; Bollen and Curran 2006).

Unconditional LGM for binary growth outcomes

As described previously, one of the assumption made in formulating LGM from the CFA model was that the outcome variable is continuous and normally distributed (Meredith and Tisak 1990). Hence, application of this modelling technique to categorical data in general (Mehta, Neale et al. 2004) and to binary variables in particular (Bollen and Curran 2006) is based on the assumption that there is a continuous normally distributed outcome variable y_{it}^{*} corresponding to the measured outcome variable under investigation y_{it} where i corresponds to an individual subject whose growth is measured at time t. There are a number of situations where the response of interest for each study participant is recorded on a binary scale (i.e. yes/on response for each study participant) and the relationship of this binary variable with growth over time is of particular interest. For example, it is meaningful to assume nutritional status of an infant evaluated at six months of age as a continuous variable and a variable that labels each infant as being stunted or not being stunted at the age of six months as an observed binary variable.

When the outcome of interest is recorded on a binary scale the unconditional LGM described for continuous outcome variables needs modification in terms of its model specification and moment structure hypotheses, and an appropriate alternative estimator should be used in place of ML to fit the LGM (Mehta, Neale et al. 2004; Bollen and Curran 2006). In cases of a continuous growth outcome variable y_{it}, the outcome is assumed to be the same as the underlying continuous outcome variable y_{it}^{*} (i.e. $y_{it}^{*} = y_{it}$). However, the observed binary outcome variable y_{it} is the collapsed version of the underlying continuous outcome variable y_{it}^{*} and hence the same assumption of equality of

the two does not hold (i.e. $y_{it}^{*} \neq y_{it}$). As a result of this inequality, the unconditional LGM described for continuous growth outcome variable, say $y_{it} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it}$, is also true for the underlying continuous variable y_{it}^{*} but it does not hold for the observed binary outcome variable y_{it} (i.e. $y_{it}^{*} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it}$ and $y_{it} \neq \alpha_i + \lambda_t \beta_i + \varepsilon_{it}$). This discrepancy can be modified by introducing the concept of thresholds into the subject level equation as described below (Muthen and Asparouhov 2002; Mehta, Neale et al. 2004; Bollen and Curran 2006) so that the observed binary variable is mapped onto the underlying continuous latent variable.

For the observed binary variable y_{it} the threshold means the cut-off value on the underlying continuous latent variable y_{it}^{*} that dichotomizes it so that any value smaller than the threshold is mapped onto one category of the observed binary variable y_{it} and those greater than the threshold are mapped onto the remaining category of the observed binary variable. In other words, a study participant is classified as being in one or the other category of the observed variable y_{it} (e.g. stunted or not stunted, underweight or not underweight) depending on its relative position on the underlying continuous outcome variable y_{it}^{*} in reference to the threshold. In the true sense thresholds are unobserved latent values on the underlying continuous variable that have generated the observed values of the binary variable and they should be estimated.

An example from nutritional data might be helpful to clarify the meaning of threshold although the cut-off values used in practice are not estimated from the model and hence this is not necessarily equivalent to the true threshold values. Infants are classified as being stunted if their length is less than two standard deviation (sd) units from the median length of the WHO reference population of infants of the same age. In this example we are using the idea of threshold with the assumption that WHO reference population (or any other reference population) of infants is representing the underlying normally distributed nutritional status variable at a given age and gender combination. However, in this example we are just using -2*sd as a cut-off rather than estimating the threshold of the underlying distribution, although it should be estimated in the true sense.

Let τ_{1_t} be the cut-off value of the underlying continuous latent growth outcome variable y_{it}^* at time t that dichotomizes y_{it}^* into two groups so that the observed growth measurement y_{it} assumes one of its two possible values (e.g. stunted or not being stunted at six month of age) depending on which direction of τ_{1_t} an infant is on its underlying y_{it}^*. The following formulation of the unconditional LGM appropriately maps the observed binary growth measure y_{it} with the underlying continuous latent growth measure y_{it}^* through the threshold τ_{1_t}

$$\begin{cases} y_{it}^* = \alpha_i + \lambda_t \beta_i + \varepsilon_{it} \\ y_{it} = 1 \ \ if \ \ y_{it}^* \le \tau_{1_t} \ \ and \ \ y_{it} = 0 \ \ if \ \ y_{it}^* > \tau_{1_t} \end{cases}$$

In this threshold model, the mean (i.e. E(y_{it}^*), variance (i.e. var(y_{it}^*)), threshold (i.e. τ_{1_t}) and probability distribution of the underlying continuous latent growth outcome variable (i.e. y_{it}^*) are all unknown. On the other hand there is a binary observed outcome variable which will result in the observed proportion of study participants who are classified in the category of interest (e.g. the proportion of stunted infants at the age of six months). Having a model with four unknown parameters and one known summary from the data implies an under-identified model (Bollen 1989).

Making the distributional assumption that the underlying continuous latent growth variable has a normal distribution reduces the unknown parameters to be estimated to three (i.e. mean, variance and threshold of the normally distributed variable) and one proportion to be estimated from the observed binary data, which is still an under-identified model. Further assumption of the standard normal distribution for the underlying variable makes the model just identified and the only unknown parameter will be the threshold. Hence, the threshold would be the same as the probit (i.e. the inverse of the cumulative probability of a standard normal distribution) of the observed proportion of subjects that falls in the lower category of the observed binary outcome variable.

In LGM of a binary outcome variable (e.g. stunting of infant over the first eighteen months) unless the restrictive assumption of the standard normal distribution is used which guarantees a mean of zero and variance of one model identification requires fixing two of the three parameters (mean, variance, and threshold) of the underlying continuous

latent growth variable so that the third parameter can then be estimated (Bollen and Curran 2006). For example, while modelling change of stunting and underweight of infants over time it is meaningful to monitor the trend in the prevalence of undernutrition (i.e. estimate the mean of the underlying latent growth variables). This can be achieved by fixing the threshold and the variance of the underlying nutritional status of infants.

The second problem in applying the LGM technique developed for continuous outcome variables to model the observed binary growth variables is related to the moment structure hypotheses. In modelling growth of continuous outcome variables the moment structure hypotheses are stated as (Bollen and Curran 2006):

$$\sum = \sum(\theta) \quad and \quad \mu = \mu(\theta).$$

In this situation the input data for LGM are vectors of means and covariances of continuous variables. The hypothesized model is expected to describe the pattern of the observed vector of means and covariances in terms of a few estimated model parameters.

For the observed binary outcome variables, say y, whose underlying continuous latent variable is denoted by y^*, moment structure hypotheses are stated in terms of the underlying variable (Bollen and Curran 2006) as follows:

$$\sum{}^* = \sum(\theta) \quad and \quad \mu^* = \mu(\theta)$$

where

- $\sum{}^*$ is the population covariance matrix of y^*,
- μ^* is the vector of means of y^*,
- $\sum(\theta)$ is the model implied covariance matrix,
- $\mu(\theta)$ is the model implied mean vector, and
- θ is the vector of model parameters.

Relevant summary statistics from the observed binary outcome variables that can be used as an input for the LGM to test the above moment structure hypotheses are (1) marginal proportions for the categories of each variable and (2) proportions within the cells of multidimensional contingency tables showing the frequencies of the patterns with which different categories of the outcome variables occur. For example, if infants are classified

as being stunted or not at the ages of two, six, nine, 12, and18 months, the data that will be used in the LGM will be (1) the proportion of stunted infants at each of the follow-up time points, and (2) the proportion of stunted infants in each cell of the contingency table formed by cross-tabulation of the five follow-up time points. The hypothesized model is expected to predict the proportion of individuals in the categories of each outcome variable and the proportions of individuals within the cells of the multidimensional contingency table in terms of hypothesized model parameters (Mehta, Neale et al. 2004).

If we extract means and covariances, which are necessary to define LGM, of the underlying normally distributed latent variables from the summary statistics of the observed binary outcome variables (i.e. marginal and cell proportions of the contingency table) the hypothesised moment structure hypotheses will hold for the underlying variable (Mehta, Neale et al. 2004; Bollen and Curran 2006). In other words if we are able to extract μ^* and $\sum^*$ for the underlying continuous latent variable y* from the observed marginal proportions of the observed binary outcome variables and cell proportions of the contingency table then the moment structure hypotheses $\sum^* = \sum(\theta) \quad and \quad \mu^* = \mu(\theta)$ will hold for y*. However, the true marginal and joint proportions of binary outcome variables will not be the same as the extracted mean vector and covariance matrix of the underlying latent variables (i.e. $\sum \neq \sum^*$ and $\mu \neq \mu^*$). This implies that the matrix of joint probabilities of the binary outcome variables and vector of marginal proportions of binary outcome variables will not be equal to their corresponding model implied mean vector and variance-covariance matrices (i.e. $\sum \neq \sum(\theta) \quad and \quad \mu \neq \mu(\theta)$). This means that the moment structure hypotheses will typically not hold for the observed repeated binary measures, even when they do hold for the underlying continuous repeated variables (Mehta, Neale et al. 2004; Bollen and Curran 2006).

If it is reasonable to assume a bivariate normal distribution for the underlying latent variables (e.g. stunting of infants at any two follow-up time points) tetrachoric correlation is an appropriate measure of the degree of association between two binary variables (Olsson 1979). Hence, some of the problems related to the moment structure hypotheses can be addressed by estimating mean vector and diagonal elements of the

covariance matrix (i.e. variance vector) using the threshold model (Olsson 1979; Bollen and Curran 2006). Similarly, the problem associated with the off-diagonal elements of the variance covariance matrix (i.e. covariance) is addressed by estimating tetrachoric covariance from the sample data in two steps (Olsson 1979; Bollen and Curran 2006). In the first step, tetrachoric correlation between each pair of underlying continuous latent variables (e.g. stunting at two months and stunting at six months) are estimated. The matrix of tetrachoric correlation (i.e. $\hat{R}^*$) is then assembled from these pairwise estimates. In the second step the tetrachoric covariance matrix is calculated (Olsson 1979; Bollen and Curran 2006) as follows:

$$\hat{\Sigma}^* = D\hat{R}^*D$$

where

- $\hat{\Sigma}^*$ is estimated tetrachoric covariance matrix,
- D is a diagonal matrix of the standard deviations of the underlying repeated measures estimated from the threshold model,
- $\hat{R}^*$ is a matrix assembled from the pairwise estimation of tetrachoric correlations.

After extracting the mean vector and the variance-covariance matrix for the underlying normally distributed latent variable as described above, (a) from the marginal proportions and (b) from the proportions of the joint cells of the multidimensional contingency table constructed from the observed binary variables, parameters of the LGM can be estimated using ML or WLS estimators (Muthen and Muthen 1998-2009; Bollen and Curran 2006).

Conditional Latent Growth Model

After attaining a satisfactory unconditional LGM for the repeated measures regardless of the scale of the outcome variable (i.e. continuous or categorical), a conditional LGM can be fitted with the objective of obtaining an explanation for the variability that might have been identified in the parameters of individual level LGM (Curran and Hussong 2003; Biesanz, Deeb-Sossa et al. 2004; Bollen and Curran 2006). For example, if the best fitting unconditional LGM for the repeated measure has two parameters, say intercept factor α_i and slope factor β_i , with significant variability among individual study participants (i.e. σ_α^2 and σ_β^2 are significantly different from zero) and if we assume that

there are two hypothesized characteristics, say x_1 and x_2, that could potentially explain the variability observed within these growth trajectory parameters, then the conditional LGM can be specified as follows(Curran and Hussong 2003):

$$\alpha_i = \mu_\alpha + \gamma_1 x_{i1} + \gamma_2 x_{i2} + \zeta_{\alpha_i}$$
$$\beta_i = \mu_\beta + \gamma_3 x_{i1} + \gamma_4 x_{i2} + \zeta_{\beta_i}$$

where

- ζ_{α_i} and ζ_{β_i} are discrepancies between the i^{th} individual trajectory intercept and slope, respectively, which could not be explained by the two covariates
- μ_α is the trajectory intercept and μ_β is the trajectory slope of an average study participant adjusted for the effects x_1 and x_2
- $\gamma_1, \gamma_2, \gamma_3, \gamma_4$ are the fixed effects of the corresponding covariate on the parameter of the trajectory of the i^{th} study participant. The values of the $\gamma's$ remains the same from individual to individual and are known as fixed effects. The effect of x_1 and x_2 on the trajectory parameters will only depend on the individual specific values of x_1 and x_2.

In the practical application of LGM the two equations presented above to predict model parameters are substituted into the best fitting unconditional LGM resulting in conditional LGM, and all model parameters are estimated simultaneously. For a continuous outcome measured on more than one occasion, the best fitting unconditional LGM with two parameters described before and the two covariates (x_1 and x_2) as predictors of the two trajectory parameters, the full conditional LGM is

$$\begin{aligned} y_{it} &= \mu_\alpha + \gamma_1 x_{i1} + \gamma_2 x_{i2} + \zeta_{\alpha_i} + (\mu_\beta + \gamma_3 x_{i1} + \gamma_4 x_{i2} + \zeta_{\beta_i})\lambda_t + \varepsilon_{it} \\ &= (\mu_\alpha + \mu_\beta \lambda_t) + (\gamma_1 x_{i1} + \gamma_2 x_{i2}) + \gamma_3 \lambda_t x_{i1} + \gamma_4 \lambda_t x_{i2} + (\zeta_{\alpha_i} + \zeta_{\beta_i} \lambda_t + \varepsilon_{it}) \end{aligned}$$

This conditional LGM assumes that both x_1 and x_2 explain some of the variability in the intercept and the rate of change of individual trajectories estimated using unconditional LGM, and there are three sources of variability (two time-dependent and one time-independent) that might need to be explained by other covariates:

- γ_1 and γ_2 are measures of the fixed effects of x_1 and x_2, respectively, on the intercept of individual trajectories

- γ_3 and γ_3 are measures of the fixed effects of x_1 and x_2, respectively, on the rate of change of individual trajectories. To evaluate whether a covariate has significant effect on the individual rate of change or not the interaction term of that covariate and measure of time (eg. $\lambda_t * x_{i1}$) is include in the model and its coefficient (eg. γ_3) is estimated.
- ζ_{α_i} is the deviation between the i^{th} individual trajectory intercept and the intercept of the trajectory of an average study participant. This deviation can be reduced by including factors that can potentially affect the intercept of the trajectory as a predictor of the intercept in the conditional LGM
- ζ_{β_i} is the time specific deviation of i^{th} individual rate of change and the rate of change of an average study participant. This deviation can be reduced by including covariates (i.e. interaction of time and factors that can potentially affect the rate of change of individual trajectory) within the conditional LGM
- ε_{it} is random variability of the outcome of interest (e.g. y_{it}) for the i^{th} study participants on the measurement time t.

The above conditional LGM also assumed that the two covariates (x_1 and x_2) are independent predictors of the trajectory parameters. If there is an *a priori* hypothesis that one of the two covariates (e.g. x_2) is a mediator of the association between the remaining covariate (e.g. x_1) and the trajectory parameters (intercept or rate of change or both) the unconditional LGM presented above can be modified in order to take account of this situation. However, the formulation of the model is easier using matrix notation or path diagrams. The method of path diagram will be described in the methods chapter of this thesis in connection with the hypothesised mediating role of birth weight and other factors in evaluating the effect of antenatal CMD on infant growth.

1.7.1.3 Evaluating Model fit in SEM

In structural equation modelling, the main interest is to find a meaningful explanation for the association of variables simultaneously. This association can be analysed using the raw data or variance and covariance matrix which has sufficient information about the association. In the case of LGM, the vector of means is also part of the input used in

analyzing associations and evaluating goodness-of-fit of hypothesised models. Evaluating the validity of the central null hypothesis (i.e. $\sum = \sum(\theta)$) in SEM involves deciding how closely the population variance and covariance matrix is reproduced by the variance and covariance matrix implied by the model under investigation, with the hope of explaining the target variation in terms of the parameters of this model (Bollen 1989). If we include comparison of mean values as well as the covariance structure (for example, the case of LGM) the central null hypothesis and its evolution focuses on both mean and variance and covariance structures simultaneously (Bollen and Curran 2006). In other words the central null hypotheses for the LGM is stated as $\sum = \sum(\theta)$ *and* $\mu = \mu(\theta)$ where μ is the population mean vector and $\mu(\theta)$ model implied mean vector expressed in terms of model parameters.

To evaluate the null hypothesis discrepancy functions $f_{fit} = f(S, \hat{\Sigma}(\theta))$ for SEM without mean structure and $f_{fit} = f(S, \bar{X}; \hat{\Sigma}(\theta); \hat{\mu}(\theta))$ for LGM (f_{fit} are also called fitting functions) are defined as measures of closeness, where

- S is an unbiased estimate of population variance and covariance matrix obtained from sample data,
- $\hat{\Sigma}(\theta)$ is the model implied variance and covariance matrix,
- $\bar{X}$ is an unbiased estimate of the population mean vector,
- $\mu(\theta)$ is the model implied mean vector, and
- f_{fit} is minimized with respect to θ to generate model parameters (Bollen 1989; Bollen and Curran 2006).

The Chi-square distribution has a long history serving as means of evaluating goodness-of-fit in the structural equation modelling literature (Anderson and Gerbing 1984). The statistic $\chi^2 = (N-1) * f_{\min}$ where N stands for total number of subjects and f_{min} stands for the minimum of the fitting function f_{fit} in the domain of the parameter space, has the central chi-square distribution provided that the null hypothesis is true (Bollen 1989). It is used to assess global goodness-of-fit of the hypothesized model by quantifying the magnitude of the discrepancies between the sample and model implied covariance matrices in SEM without mean structure. In assessing the goodness-of-fit of LGM it also

considers the magnitude of the discrepancies between sample mean vector and model implied mean, in addition to the covariance matrix. In other words, chi-square provides a test of whether residual differences between the sample summary (i.e.covariance and mean vector) and corresponding model implied summaries (covariance matrices and mean vector) converge in probability to zero as the sample size approaches infinity (Cudeck and Browne 1983; Bollen and Curran 2006).

Researchers using structural equation modelling routinely report the value of chi-square and associated p-value as supportive evidence to accept or reject the central null hypothesis (Hu and Bentler 1995; McDonald and Ho 2002). However, difficulties associated with its use as the goodness-of-fit have been recognized (Bentler and Bonett 1980; Hu, Bentler et al. 1992) and have provoked ongoing discussion (Bentler 1990; Kaplan 1990; Barrett 2007; Bentler 2007). The main concerns with the use of chi-square include its dependence on sample size (Bentler and Bonett 1980; Bentler 1990; Hu, Bentler et al. 1992; Marsh and Balla 1994), its dependence on normal distribution of indicator variables which might not always be attainable with practical research data, its performance under misspecification of the hypothesized model, and its dependence on the discrepancy function minimized to estimate model parameters (Hu, Bentler et al. 1992; Marsh and Balla 1994; Hu and Bentler 1995; Marsh, Hau et al. 2005).

It is well known that the power of a test increases with sample size. In goodness-of-fit tests of SEM this implies an increased probability of identifying that the model does not explain the hypothesised structure of the data. However, the value of chi-square can always be made larger by increasing sample size and almost every proposed model will be rejected if the sample size is large regardless of the magnitude of discrepancy (Bentler and Bonett 1980; Bentler 1990; Marsh and Balla 1994). The numerical value of chi-square and associated p-value are also not helpful in differentiating between statistical significance and degree of fit of the model (Gerbing and Andersen 1993).

Moreover, the Chi-square statistic defined as a function of $f_{\min}$ might not always be appropriate to use as a goodness-of-fit measure because of the following reasons (Bentler 1990): (a) in cases where some or all the assumptions underlying the fit function are violated the test might not be robust to these violations, (b) in studies involving small

sample sizes the fit function might not follow the claimed chi-square distribution and, (c) in studies involving very large sample size any trivial discrepancy might become statistically significant supporting rejection of the null hypothesis. To overcome these and associated reasons several attempts were made to obtain alternative measures by transforming goodness-of-fit tests based on $f_{\min}$ into more interpretable indices whose value lie in the 0 to 1 range (Bentler 1990).

One of the key underlying assumptions for $\chi^2 = (N-1) * f_{\min}$ to follow central chi-square and to be used as a measure of goodness-of-fit is that there is no model misspecification in the hypothesised model and that the hypothesized model can reproduce the population variance-covariance matrix of the measured variables without any error (Bollen 1989; Curran, Bollen et al. 2002). However, models specified in research are always approximations of true population models since there will be always misspecifications of unknown magnitude (Cudeck and Browne 1983). As a consequence of this, if model misspecification is moderate, there is no excessive kurtosis and a moderately large sample size, the appropriate asymptotic distribution of the statistic $f_{\min}$ and hence $\chi^2 = (N-1) * f_{\min}$ is not chi-square but it is the non-central chi-square defined by its degrees of freedom and non-centrality parameter δ (Steiger and Lind 1980; Curran, Bollen et al. 2002).

The non-centrality parameter, δ, which characterizes the distribution of non-central chi-square measures degree of misfit or the shift in chi-square distribution if the alternative hypothesis is true compared to the location of chi-square distribution we would expect if the null hypothesis was true. Mean and variance of central chi-square distribution are equal to the degrees of freedom (df) and 2*df, respectively, but the mean and variance of non-central chi-square distribution are $df + \delta$ and $2 * df + 4\delta$, respectively. Researchers have used this non-centrality parameter and associated confidence intervals as the basis for evaluating the degree of model misfit to overcome some of the difficulties associated with chi-square and p-values (Steiger and Lind 1980; Steiger 1990; Curran, Bollen et al. 2002).

Another concern regarding the use of any statistic based on the fit function f_{fit} as a measure of goodness-of-fit is its varying performance depending on the fit function used at a given time (Anderson and Gerbing 1984). There are several alternative discrepancy functions in SEM and some of them rely on more restrictive assumptions than others (Bollen 1989). Different parts of one SEM might not necessarily behave in the same way under various fitting functions (Benson and Fleishman 1994) and a single overall goodness-of-fit measure might not be capable of responding to each of these component misfits.

To supplement chi-square in model evaluation several fit indices have been developed (Hu and Bentler 1995; Marsh, Hau et al. 2005) starting in the early 1970s (Tucker and Lewis 1973) and their performance has been studied extensively (Bearden, Sharma et al. 1982; Marsh, Balla et al. 1988; Gerbing and Andersen 1993; Hu and Bentler 1998). Appreciating the large sample effect on the log-likelihood ratio as the test statistic, the Tucker-Lewis index (TLI) which is defined as follows is the oldest fit index (Tucker and Lewis 1973):

$$TLI \doteq \frac{\left(\chi_o^2/df_0\right)-\left(\chi_k^2/df_k\right)}{\left(\chi_o^2/df_0\right)-1} = \frac{Q_0 - Q_k}{Q_0 - 1}$$

where

- χ_0^2 is the likelihood chi-square for the baseline model
- χ_k^2 is the likelihood chi-square for the hypothesized model
- df_0 is the degrees of freedom of the baseline model
- df_0 is the degrees of freedom of the hypothesized model

This index was primarily derived as a measure of reliability to help in deciding on the number of factors to retain in EFA and its derivation was based on the analogy of variance component of the analysis of variance where (1) there are n observations recorded on p variables, (2) it is planned to extract k common factors to summarize the variance-covariance matrix using the maximum likelihood factor analysis method and, (3) a complete independence of variables or absence of any common factor is assumed as a baseline comparison model (Tucker and Lewis 1973). As a measure of reliability the numerical value of TLI has an interpretation as the amount of increase in fit obtained by

using k common factors compared to the no common factor situation (Tucker and Lewis 1973; Bentler and Bonett 1980).

Seven years later two comparative indices whose possible value lies between zero and one, namely, non-normed fit index (NNFI) and normed fit index (NFI) were introduced (Bentler and Bonett 1980):

(a) non-normed fit index, $NNFI = \delta_{kl} \doteq \dfrac{\left(\chi_k^2/df_k\right)-\left(\chi_l^2/df_l\right)}{\left(\chi_0^2/df_0\right)-1} = \dfrac{Q_k - Q_l}{Q_0 - 1}$

This index is the generalization of TLI such that it can be used to compare the fit of hierarchical models under different discrepancy functions and different estimation methods. In defining NNFI, three different models are involved leading to generation of three likelihood chi-square values with their corresponding degrees of freedoms, namely, two hypothesised hierarchical models and the null model of independence.

(b) a more general normed fit index : $NFI = \dfrac{f_k - f_l}{f_o}$ where f is any fit function, f_k and f_l are the minimum of the two hypothesized hierarchical models, and f_o is the fit function evaluated under the null model of independence.

A value of 0.9 and higher have been recommended to be used as indicator of best fitting models for these and any comparative fit index whose value lies between 0 and 1 inclusive. The values of these $f's$ are of the form $f_0 > f_k > f_l > 0$ and possible values of the new fit index are between zero and one inclusive.

Parallel to the newly developed comparative fit indices a different approach of model evaluation was considered and a new index whose possible value ranges between zero and one was introduced into the SEM literature based on the idea of quantifying how badly the proposed model fits to the data (Steiger and Lind 1980). This is now referred as standardized root mean square error of approximation (SRMEA) (Steiger 1990). The authors are also credited with introducing a method for obtaining confidence intervals around the newly developed non-centrality-based index using the non-central chi-square distribution giving a different dimension to the hypothesis testing tradition of SEM. The RMSEA formula is (Bollen and Curran 2006):

$$RMSEA = \sqrt{\frac{\chi_k^2 - df_k}{(N-1) * df_k}}$$

where

- χ_k^2 is the likelihood chi-square from the target model and df_k is its associated digress of freedom
- the quantity $\chi_k^2 - df_k$ is an asymptotically unbiased estimator of the noncentrality parameter for the non-central chi-square distribution underlying χ_k^2,
- N-1 in the denominator is used for adjusting the effect of sample size on the non-centrality parameter, and
- df_k in the denominator is meant to provide a penalty for using model degrees of freedom.

A recent review of fit indices showed that invention of new indices as well as re-invention of the existing ones under different names has significantly increased to the extent of making selection among them difficult and inconstant across studies (Marsh, Hau et al. 2005). To overcome these difficulties, the performance of several fit-indices were evaluated using Monte Carlo studies (Bearden, Sharma et al. 1982; Anderson and Gerbing 1984; Marsh, Balla et al. 1988; Hu and Bentler 1998) and effort was made to classify them into functional families (Marsh, Balla et al. 1988; Hoyle and Panter 1995; Hu and Bentler 1998; Fan, Wang et al. 1999; Hu and Bentler 1999; McDonald and Ho 2002; Marsh, Hau et al. 2005).

Classification of these indices into absolute and incremental fit is now becoming common practice (Marsh, Balla et al. 1988; Bollen 1989; Gerbing and Andersen 1993; Hu and Bentler 1995; Hu and Bentler 1998; Hu and Bentler 1999). Absolute (or standalone) indices assess how much the hypothesized model reproduces the population covariance matrix and they only use information from the target model. However, incremental fit indices (also called comparative fit indices) assess amount of improvement gained from using the target model compared to the baseline model. Depending on the assumption they make about the fit function and the sample data, incremental fit indices are further divided into type I, type II and type III. Of these indices type I uses information only from fit functions without making any particular

distributional assumption about the fit function, type II uses the same information as type I but assumes the central chi-square distribution for the fit function and also uses the expected value of the fit function, and type III uses the same information as that of type II incremental fit indices but under the assumption of the non-central chi-square distribution for the fit function.

In the presence of all these indices, which might not always support the same conclusion in a given data set and hypothesized model, it is not easy to make selection in a principled manner (Hu and Bentler 1995; MacCallum and Austin 2000; McDonald and Ho 2002) which makes comparison of findings across studies difficult. Recommendations are available on how and what to report from the result of SEM (Hoyle and Panter 1995; McDonald and Ho 2002; Marsh, Hau et al. 2005). The current recommendation is to report at least one index from each family (Marsh, Balla et al. 1988; Hu and Bentler 1998; Fan, Wang et al. 1999; Hu and Bentler 1999; McDonald and Ho 2002; Marsh, Hau et al. 2005). Computational software also tends to produce fit indices stratified by recommended families (e.g. AMOS version 7) or provide only selected fit indices from each family (Muthen and Muthen 1998-2009)

Besides the selection of best performing fit indices some work has also been done on their cut-off values. Recommendation of 0.9 as a measure of good fit (Bentler and Bonett 1980) was almost universally accepted for normed fit indices for the first few years. More recently an influential recommendation has been made for a cut-off value of 0.95 (Hu and Bentler 1999) based on a large Monte Carlo study (Hu and Bentler 1998), and this is now being used extensively in applied research (McDonald and Ho 2002; Marsh, Hau et al. 2005). Regarding the root mean square error of approximation (RMSEA) (Steiger and Lind 1980) cut-off values of 0.05 and 0.08 have been recommended (Browne and Cudeck 1993) to indicate the close fit and a reasonable error of approximation, respectively, until a recent recommendation of 0.06 to replace 0.05 as an upper bound for close fit (Hu and Bentler 1998; Hu and Bentler 1999).

Regardless of the continuous effort to find convincing evidence on how to use fit indices efficiently, a significant increase in the number of publications reported using various fit indices and the extensive usage of recommended cut-off values in making decisions about model fit, there are still controversies around them (Marsh, Hau et al. 2005; Barrett

2007; Bentler 2007). This controversy has reached the level of giving the recommendation not to use them at all (Barrett 2007), although others still recommend further studies (Fabrigar, Wegener et al. 1999) while using them appropriately (Bentler 2007; Steiger 2007)

In this thesis I am going to use Chi-square and associated p-value, RMSEA (or RMR), TLI and CFI to evaluate fit of models. This choice is inline with current practice and recommendations (Marsh, Hau et al. 2005) as well as their availability in available software (Arbuckle 1995-2006; Muthen and Muthen 1998-2009). These indices are relatively less sensitive to sample size and robust to non-normality (Hu and Bentler 1999) although further studies have been recommended to evaluate the response of all fit-indices to model misspecifications (Fan and Sivo 2007).

1.7.2 Multilevel Growth modelling (MGM)

1.7.2.1 General introduction to multilevel modelling

Multilevel modelling is a technique that expands generalized linear models to accommodate the situation of clustering of responses, so that the data are permitted to exhibit correlated and non-constant variability (Bryk and Raudenbush 1987; Raudenbush and Bryk 2002). It recognises the existence of clusters or hierarchies within the data by allowing residual components at each level in the hierarchy. Unlike generalized linear models, which are designed to model the means of the data allowing residuals only at the subject level, multilevel modelling gives the flexibility of modelling the means of the data as well as their variances and covariance. The objective in using multilevel modelling is, therefore, to design a model that takes advantage of correlation among responses belonging to the same cluster while modelling the association of an outcome and its hypothesized correlates (Raudenbush and Bryk 2002).

Use of multilevel modelling (Rogosa and Willett 1985; Bryk and Raudenbush 1987; Raudenbush and Bryk 2002) provides an opportunity to investigate two main research questions simultaneously which couldn't be answered using generalized linear models: (a) is there a significant difference in the response of individual subjects within a cluster? – within cluster variability, and (b) is there significant variability between response of

subjects that constitute different clusters? – between-cluster differences. For example, a two-level model which allows for grouping of child outcomes within household would include residuals at the child and household level. Thus the residual variance is partitioned into a between-household component (the variance of the household-level residuals) and a within-household component (the variance of the child-level residuals). The household residuals, often called 'household effects', represent unobserved household characteristics that affect child outcomes. It is these unobserved variables which lead to correlation between outcomes for children from the same household.

In a hypothetical example of vaccination coverage of under five children it is possible to define one binary outcome variable, for example, whether a given child has completed required vaccinations before the age of five years, and to collect information from all children in a district. In this example there will be clustering within the household that creates dependence among the vaccination status data that will be obtained from children within households. Vaccination of a child can be affected by several factors including factors specific to the child (e.g. child gender, child health, place of delivery) and household level factors (e.g. socio-economic status, parental education, family size). For a child level factor (x_1), a household level factor (x_2), and a probability of child i in household j being fully immunized at the time of the survey (π_{ij}), we can formulate a two level logit model which allows residuals to appear at household level and child level):

$$\log it(\pi_{ij}) = \beta_0 + \beta_1 x_{1i} + \beta_2 x_{2ij} + u_i + \varepsilon_{ij} \quad , \; i = 0, 1, \ldots n_j, \; j = 1, 2, \ldots N$$

where

- N is number of households considered in that investigation,
- n_j is number of children in j^{th} household,
- u_i is a household level random error generated from the normal distribution and independent from child level random error ε_{ij},
- ε_{ij} is child level independent residual randomly generated from logistic distribution,
- β_0 is the intercept,
- β_1 is the slope of the household level covariate
- β_2 is the slope for child level covariate.

Except for the additional random component u_i the above model is the same as the ordinary logit model. The random component u_i represents the effect caused by household i and it is assumed to be normally distributed with mean zero and variance σ_u^2, $N(0,\sigma_u^2)$, provided that the households are random samples from a large population of households. The inclusion of the new random term u_i in the ordinary logit model formulation is what makes the model different and it enable us to estimate the amount of household level variability that can be explained from the total variability of the outcome variable.

Multilevel modelling have been used in different disciplines but the literature appears under different names including multilevel linear models in sociological research, mixed-effects models or random-effects models in biometric applications, random-coefficient regression models in econometrics and covariance components models in the statistical literature (Raudenbush and Bryk 2002). In the field of education this method of data analysis is preferable over traditional modelling techniques to evaluate educational performance where clustering of students within sections and sections within schools are a natural phenomenon. As the result of methodological advances, researchers in different disciplines have benefited from this modelling technique to adjust for clustering of response which occurs due to hierarchal sampling techniques, cluster sampling or repeated measurement of individual subjects over time. Duncan and his colleagues (Duncan, Jones et al. 1998) have used a simple approach to popularize multilevel modelling over traditional modelling techniques in health research and Leyand and colleagues (Leyland and Groenewegen 2003) have made a clear case to show the relevance of multilevel modelling to public health policy .

In a situation where there is only one level of clustering (e.g. infant length measured over time or immunization status of under five children within households in a given community) there is a related method to multilevel modelling called generalized estimating equations (GEE) (Twisk 2003). This method can be used to adjust for the clustering effect while modelling the outcome of interest. The GEE method defines the covariance structure of observations within a cluster and incorporates this covariance structure into the model estimation procedure in a way of extending generalized linear

models (Twisk 2003). In contrast, the multilevel modelling technique defines a separate set of model parameters that accounts for the correlation within each cluster and estimates them simultaneous with other model parameters. In other words the GEE methods adjusts for the clustering effect and estimates population averaged effects but multilevel modelling specifically estimates parameters that account for the within-cluster correlation with other parameters of the model (Hu, Goldberg et al. 1998; Twisk 2003).

1.7.2.2 Multilevel growth modelling for longitudinal studies

Multilevel modelling in longitudinal studies extends generalized linear models in such a way that the interdependency of observations created due to repeated measurement of the same individuals over time are analysed while random errors are still assumed to be independent (Bryk and Raudenbush 1987; Raudenbush and Bryk 2002; Twisk 2003). For a continuous outcome variable measured over time, say length of infants, use of ordinary linear regression analysis with the assumption of a linear increase in length over time would mean to fit the following model:

$$length = \beta_0 + \beta_1 * time + \varepsilon$$

$$\varepsilon_{ij} \sim IND(0, \sigma^2)$$

where

- length = measurement of length of an individual infant at a given time point
- β_0 and β_1 are the usual regression coefficients
- ε is random error for a given infant at a given time point

If a polynomial function of a different degree is hypothesised for an increase of length of infants over time, the above model can be easily extended by including relevant polynomial terms of time.

For the validity of linear and logistic regression, and other generalized linear models, independence of residuals is one of the key assumptions that should be satisfied. In longitudinal studies individual subjects are measured repeatedly over time and these observations are naturally correlated violating the assumption of independence. In the P-MaMiE study (Hanlon, Medhin et al. 2009), for example, physical growth of infants was measured at two, six, nine, twelve and eighteen months of age and these measurements are correlated within each infant. Modelling the length of these infants over time using

linear regression would mean fitting an ordinary regression model with one outcome(length) and one independent variable (time of measurement). However, the assumption of independence of residual errors is not valid and fitting the above linear regression model would not be correct.

One possible way of correcting the above linear regression model is to include the infant identifier as a covariate and fit the following regression model, where the regression coefficients and residual errors have the same interpretation and assumptions respectively:

$$length = \beta_0 + \beta_1 time + \left(\sum_{k=1}^{p-1} \beta_k * child_k\right) + \varepsilon$$

This regression model assumes that length of an individual infant increases linearly over time and it allow all infants to have one common slope β_1 (rate of growth over time) and every infant to have a separate intercept $\beta_0 + \beta_k$, $k = 1, 2, ...p-1$ and β_0 for infant p implying as many intercepts as the number of infants in the data set.

Although the inclusion of an infant specific intercept in the regression model would be expected to reduce correlation of residuals within individuals over time the degree of reduction might not always be to the required level. Further reduction of correlations among residuals can be attained by relaxing the assumption of common slope in the above regression model allowing separate rate of growth for each infant. This can be attained by including interaction terms of time and child identification number within the above regression model:

$$length = \beta_0 + \beta_1 time + \left(\sum_{k=1}^{p-1} \beta_k * child_k\right) + \left(\sum_{j=1}^{p-1} \beta_j time * child_j\right) + \varepsilon$$

This regression model assumes linear increase of infant length over time and allows each infant to have a separate intercept and separate slope needing estimation of an extremely large number of parameters. However, this might not always be practical with the given data set.

The third option is to let all infants to have the same intercept and allow separate slope for each infant. With the assumption of linear increase of length over time this can be attained by including an interaction term in the model and excluding the infant identifier variable from the model:

$$length = \beta_0 + \beta_1 time + \left(\sum_{j=1}^{p-1} \beta_j time * child_j \right) + \varepsilon$$

A common characteristic of these three options is estimation of large number of parameters resulting in a non-parsimonious model.

Multilevel modelling takes a different approach with the objective of addressing the same problem but with a parsimonious model. Instead of estimating large numbers of regression coefficients, the focus is placed on the variability among individual regression coefficients and their co-variation (i.e. variability among individual intercepts and slopes, and covariance of these two sets of regression coefficients). In a condition where there are n study participants this approach allow us to estimate three parameters (i.e. variance of random intercept, variance of random slope and covariance of random slope and random intercept) instead of 2*n regression parameters (one intercept and one slope for each participant) making the model more parsimonious without compromising the main objective of reducing interdependence of residuals within individual infants. The approach to the above problem in multilevel modelling would be addressed by fitting one of the following three regression models:

$length_{it} = \beta_0 + \beta_1 time_t + u_{1i} + \varepsilon_{it}$ ---------- random intercept model

$length_{it} = \beta_0 + \beta_1 time_t + u_{2i} + \varepsilon_{it}$ -------- random slope model

$length_{it} = \beta_0 + \beta_1 time_t + u_{1i} + u_{2i} + \varepsilon_{it}$ ---- random coefficient model

where

- $\beta_0, \beta_1, and\ \varepsilon_{it}$ have the usual meaning and satisfy usual distributional assumptions
- u_{1i} and u_{2i} are random intercept and random slope
- $u_{1i} \sim N(0, \psi_1)$; $u_{2i} \sim N(0, \psi_2)$; $\text{cov}(u_{1i}, u_{2i}) = \omega$, $u_{1i}\ and\ u_{2i}$ are independent of ε_{it}

If linear change over time does not fit the data, a polynomial function of higher degree should be investigated before proceeding to the next step. Once the parameters of the basic multilevel growth model are estimated and their individual statistical significance is determined, the following step is to check if there is any extra variability of the outcome variable which needs to be explained. In the presence of significant variability additional covariates can be incorporated into the basic growth mode as follows:

$$length = f_1(time) + \sum_{i=1}^{k} \beta_i x_i \ , i = 1, 2, ...k$$

where

- f_1(time) is the basic multilevel growth model , and
- xi, i=1, 2, ..k are covariates expected to explain extra variability in length of infants over the first 18 months of life which is not explained by f_1(time).

If there is any covariate which might be hypothesised to influence the rate of change over time, say z, then the interaction of this variable with time (i.e. time*z) should be include as one of the covariates. However, any interaction term should be include in the model as long as the main effect is also included in the model regardless of its statistical significance.

1.7.2.3 Assessing goodness of fit of multilevel growth models

Unlike the SEM framework there are no fit indices that are recommended to assess goodness-of-fit of MGM. Like any other ordinary modelling technique, statistical significance of individual regression coefficients of the covariates are assessed using a Wald test and confidence intervals are constructed around point estimates using normal distribution assumptions. For nested models, the contribution of individual covariates or groups of extra covariates in explaining extra variability of an outcome can be evaluated using a likelihood ratio test as long as the ML method is used to estimate the model. In case of comparison between the model fit of non-nested models, Akaike Information Criterion (AIC) and the Bayesian Information Criterion (BIC) are the candidate global fit measurers, and the model with the smaller value is preferable (Singer and Willett 2003).

CHAPTER 2: AIMS, OBJECTIVES AND HYPOTHESES

2.1 Aim of the study

The primary aim of this study was to compare traditional methods of modelling binary outcome variables (i.e. logistic regression) and normally distributed continuous outcome variables (i.e. linear regression) with currently available complex modelling techniques for repeated measures (i.e. multilevel growth modelling and latent growth modelling) as applied data analysis methods. To address the stated aim we have used infant growth data from the birth cohort established in the Butajira area, Ethiopia, between July 2005 and February 2006 to evaluate the effect of perinatal maternal common mental disorders (CMD) on infant outcomes.

In this thesis we used maternal perinatal CMD as the main exposure variable. Similarly, we have used three groups of outcome variables, namely, (1) two measured infant anthropometric growth indicators (i.e. weight in kilograms and length in centimetres), (2) two derived continuous nutritional indicators in reference to the 2006 WHO reference population (i.e. weight-for-age and length-for-age z scores), and (3) two derived binary variables as nutritional indicators of infants in reference to the 2006 WHO reference population (i.e. underweight defined as weight-for-age z score < -2 and stunting defined as length-for-age z score < -2). .

2.2 Objectives to be addressed

In the process of developing the thesis using infant growth data we will evaluate

1. the effect of perinatal CMD on each of the six infant growth/nutritional outcomes (i.e. weight, length, weight-for-age z, length-for-age z, underweight and stunting) using the three different modelling techniques (traditional methods, multilevel growth modelling and latent growth modelling)

2. the roles of pre-specified locally relevant risk factors for infant growth as confounders or mediators of the association between CMD and infant growth, and as independent predictors of infant growth

3. the three modelling techniques in the light of addressing objectives 1 and 2 in terms of data requirement, simplicity, flexibility, capability, and accessibility for practical use.

2.3 Hypotheses to be evaluated

- Maternal CMD measured (a) in the third trimester of pregnancy (prevalence) or, (b) at two month postnatal (prevalence) or (c) persistent between pregnancy and two month postnatal will have a significant negative effect on physical growth (length, weight, length-for-age, weight-for-age, stunting, underweight) of infants in the first 18 months of life before and after adjusting for pre-specified risk factors

- As a result of how the modelling technique handles mediating variables perinatal CMD will show a strong negative effect on infant growth when the association is modelled using LGM rather than traditional modelling techniques and multilevel growth modelling

- Multilevel growth modelling and latent growth modelling techniques will identify similar factors affecting infant growth provided that their effects on growth were not hypothesized to be mediated by some other factors.

CHAPTER 3: METHODS

3.1 The study setting

3.1.1 Country Background

Geography

Ethiopia is located in the Horn of Africa between 3 and 15 degree north latitude and between 33 and 48 degree east longitude, with an approximate area of 1.1 million square kilometres, and bordering with Eritrea, Djibouti, Somalia, Kenya, and the Sudan (FMOH 2002; Central Intelligence Agency 2008). The country has a high central plateau that varies from 1,290 to 3,000 meters above sea level. The highest point is Ras Dashen in the northern highlands with an altitude of 4,620 meters and the other extreme, the lowest place in the world, known as the Danakil Depression, with an altitude of 125 meters below sea level is located in the north eastern part of the country. Elevation is generally highest just before the point of descent to the Great Rift Valley, which splits the plateau diagonally. The plateau gradually slopes to the lowlands of the Sudan on the west and the Somali-inhabited plains to the southeast (accessed on January 18, 2011: http://en.wikipedia.org/wiki/Ethiopia).

Climate

Rainfall and temperature patterns in Ethiopia vary widely because of its location in the tropics and its diverse topography (accessed on January 17, 2011 http://www.fas.usda.gov/pecad2/highlights/2002/10/ethiopia/baseline/Eth_Annual_Rainfall.htm). The whole country might be divided into three rainfall regions (i.e. two seasons, three seasons and four seasons) but the climate is temperate on the plateau and hot in the lowlands. Annual precipitation ranges from 800 to 2200 mm in the highlands (>1500 meters), varies from less than 200 to 800 mm in the lowlands (<1500 meters) and decreases northwards and eastwards from the high rainfall pocket area in the southwest. The western half of Ethiopia has two distinct seasons (wet from June-September and dry from November-February), with the rainfall peak occurring from July to August. The central and most of the eastern part of the country have two rainy periods known locally as the main *Kiremt* rains (June –September) and small *Belg* rains, (February-May), and one dry period known locally as the dry *Bega* season (October-January). The south and south-eastern parts of Ethiopia have two distinct dry periods (December-February and

July-August) and two rain seasons (March-June and September-November). The main rain which is referred to as the *belg* rains occurs from March-June.

Administrative Structure

At present Ethiopia follows a federal system led by the prime minister and based on the constitution ratified in 1994. This followed the fall of the previous military government that had ruled the country for 17 years (http://www.ethiopar.net accessed on 17/01/2011). There are nine administrative regions (i.e. Tigray, Afar, Amhara, Oromia, Somali, Benishangul Gumuz, Southern Nations, Nationalities and Peoples (SNNP); Gambella and Harari) and two federal cities (Addis Ababa, the capital city of Ethiopia, and Dire Dawa). Each of these 11 administrative units has three hierarchical administrative structures: zone, woreda and kebele. The woreda is the basic decentralized administrative unit. The kebele is the smallest administrative structure of the government in urban settings, reporting to the woreda, and it is synonymous with peasant association (PA) in the rural area. Across the whole country there are 80 zones, 551 woredas and about 12,000 kebeles/PAs (accessed on January 17, 2011 http://cnhde.ei.columbia.edu/healthsystem/political_admin.html) (Figure 3.1). To respond to local problems in a timely manner and facilitate development of the country, the government of Ethiopia has decentralized political power to the woreda level and the role of zone is limited to the coordination of activates. The woreda has an administrative council composed of elected members. The health delivery system to the public also follows the administrative structure of the country.

Figure 3.1: Administrative map of Ethiopia showing different zones
(Accessed on 17/01/11 http://cnhde.ei.columbia.edu/healthsystem/political_admin.html)

People

The most recent population and housing census was carried out in Ethiopia in 2007, with preliminary results was published in 2008) (Central Statistical Agency 2008). According to this report, Ethiopia has a population of 74 million, with an annual growth rate of 2.6% that varies significantly across regions, with the lowest growth rate of 1.7% in Amhara Regional State and the highest annual growth rate of 4.1% in Gambella Regional State. More than 83.0% of the Ethiopian population live in rural areas and depend on traditional agriculture for their livelihood. In the year 2010 the agricultural sector contributed 42.0% of the total national gross product (http://www.mofed.gov.et). The age structure of the population is pyramidal with the under 15 year population comprising 44% of the total, only 5% above the age of 65 years and 5% are at most one year of age. The life expectancy is 57 years for males and 60 years for females(WHO 2010).

The country is a multi-ethnic state with a multitude of languages but two ethnic groups (i.e.Oromo (34.5%) and Amhara (26.9%)) making up more than 60% of the total population (Central Statistical Agency 2008). There are more than 77 different ethnic groups with their own distinct languages within Ethiopia and some of these have as few as 10,000 members. Ethiopian Orthodox Christianity (43.5%), Islam (33.9%) and Protestant Christianity (18.6%) are the three major religions. In general, most of the Christians live in the highlands, while Muslims and adherents of traditional African religions tend to inhabit the lowland regions. Most people speak a Semitic or Cushitic language. Amharic is the official language of the country and it was the language of primary school instruction until recently. However, its role as a medium of education in elementary schools has been now replaced in many areas by local languages such as Oromifa and Tigrinya. English is the most widely spoken foreign language and is used as a medium of education in all secondary schools and universities of the country.

Structure of health systems

Ethiopia is implementing a 20-year sector program and strategy, broken into five-year rolling programs, beginning in 1997/98 (FMOH 2005). Accordingly, the first Health Sector Development Program (HSDP) took place from 1997/98 – 2001/02, the second HSDP, which was designed for a period of three years took place from 2002/03-2004/05, and the third HSDP which was recently completed took place from 2005/06-2009/10. These HSDPs addressed eight major areas or components, namely, health service delivery and quality of care, health facility construction and rehabilitation, human resources development, strengthening pharmaceutical services, Information, Education and Communication (IEC), health care financing, health management, health management information system (HMIS), and monitoring and evaluation.

The health service was also structured as a four-tier system, namely, central referral hospitals (tier 4), regional hospitals (tier 3), district hospitals (tier 2) and primary health care unit, constituted by a health centre with five satellite health posts (tier 1). The health sector strategy of Ethiopia which is used to implement the Health Policy of the country focused on giving comprehensive and integrated primary health care in health institutions with a major emphasis on community level services. Its emphasis is on preventive and promotive components, yet without neglecting the basic curative care. The main focus is

on communicable diseases, common nutritional deficiencies, and environmental health and hygiene.

Health institutions are located in the capital of the respective administrative structure of the country with the exception of health posts which are located in the central points of the peasant associations (PA) they give service to. Health centres are mainly staffed by nurses and provide outpatient and outreach clinic services both for treatment and prevention purposes. The main role of a health post is to deliver health education to the public and contribute to the prevention of diseases. People can get first aid and simple drugs from the health posts.

The Health Extension Package (HEP) was introduced in the HSDP II as a new initiative. This package is an innovative community-based health care delivery system aimed at creating a healthy environment as well as healthy living. The main objective of HEP is to improve access and equity to essential health intervention through community (*Kebele/PA*) based health services with a strong focus on sustained preventive health actions and increased health awareness. The health extension service package focuses on preventive health measures targeting households particularly women at the *Kebele/PA* level.

In rural settings every peasant association has one health post and it is staffed with two female health extension workers (HEW). This has now being extended to urban settings and female nurses with an additional six month training of the Heath Extension Packages are deployed. From the start of the program 2002/03 up to July 2010 a total of 34,384 HEW were deployed to all rural kebeles of the country and 3401 HEWs were deployed to give service in urban kebeles (FMOH 2010)

Selected development indicators

Ethiopia is one of the least developed countries in the world (UNDP 2010). It has a human development index of 0.328, attaining a rank of 157 out of 169 countries, 39.0% of its nationals live on less than $1.25 per day, 44.0% of the population live below the national poverty line, 34.6% of the under five children are underweight and healthy life expectancy at birth is 50 years (49 years in men and 51 years in women) (WHO 2010).

In 2009/2010 the percentage share of the health budget from the total national budget was 10.4% and per capita public fund allocation and expenditure on health were 39.8 birr (i.e. $2.5) and 34.55 birr (i.e. $2.2), respectively (FMOH 2010). Although Ethiopia is better off by African standards in reducing child mortality and in immunization coverage of infants, it is far behind in reducing the maternal mortality ratio and its risk factors (Table 3.1).

Health problems of the majority of Ethiopians are mainly of an infectious nature with significant variability across regions. In 2008/09, malaria, acute upper respiratory tract infection, intestinal parasitic infection and gastritis duodenitis were among the top 10 leading causes of out-patient attendance in government health institutions in the whole country. During the same year there were 13,808 hospital beds, and 45,793 health professionals (i.e. 1035 doctors, 1749 health officers, 26345 nurses, 642 pharmacy attendants and pharmacists, 717 sanitarians, 1494 health assistants, and 3667 technicians) in the government health system providing curative health services (Central Statistical Authority 2009). To address public health problems in a meaningful way the government has now focused on preventive aspect of health delivery through the Health Extension Package (FMOH 2010)

Table 3.1 selected indicators of development from performance report of the Federal Ministry of Health of Ethiopia for the year 2009/2010 and from World Health Organization Statistics 2010

Development indicators	Ethiopian situation (WHO 2010)	African average (WHO 2010)	Ethiopian situation (FMOH 2010)
Immunization coverage			
Measles	74%	73%	82.4%
DTP3	81%	72%	86.0%
HepB3	81%	67%	
Hib3	81%	38%	
Full immunization			72.3%
Neonatal mortality rate	39/1000 live birth	40/1000 live births	
Infant mortality rate	69/1000 live birth	85/1000 live births	
Under-five mortality	109/1000 live births	142/1000 live births	
Malaria mortality rate	51/100000	104/100000	
Children aged <5 years sleeping under insecticide treated nets	33%	17%	
Maternal mortality ratio	720/1000 live births	900/1000 live births	
Births attended by skilled health personnel	6%	47%	16.8%
Antenatal care coverage			
At least one visit	28%	73%	71.4%
At least four visits	12%	44%	
Postnatal care coverage			36.2%
Contraceptive prevalence	14.7%	23.7%	61.9%
Unmet need for family planning	33.8%	24.3%	
Adolescent fertility rate (i.e. among 15-19 year old girls)	109/1000 adolescent girls	118/1000 adolescent girls	
Total fertility rate	5.3/woman	4.9/woman	
Population using improved drinking water source	38%	61%	
Population using improved sanitation	12%	34%	73.7%
Gross national income per capita	$870	$2279	
Population living on less than $1.25 a day	39%	52.8%	

3.1.2 The Butajira District

Butajira district (figure 3.1) is located approximately 130 km south of the Ethiopian capital, Addis Ababa. The first modern school with two teachers was established in Butajura town in 1947 and the district acquired its present name in 1954 with its capital town Butajira (Berhane, Wall et al. 1999). Butajira town is located at an altitude of 2100m above sea level and it covers approximately 9km^2 (Berhane, Wall et al. 1999). The district is predominantly rural and residents are mainly from the Gurage ethnic group and followers of Islamic religion.

Administratively, the district belongs to the Gurage zone within the Southern Nations Nationalities and Peoples region (SNNPR). According to the most recent census results, SNNPR is the third populous region of the country with the proportional share of 20.4%, the population of Gurage zone amounts to 8.5% of the total SNNPR population and 85.5% of the population live rurally which is similar to the national average (Central Statistical Agency 2008). Due to repeated famines in the district in 1974, 1985, 1999 and 2003, the Disaster Prevention and Preparedness Commission of the Federal Government of Ethiopia has registered the area as a drought-stricken area.

With an altitude ranging from 1750 to 3400 metres, the climate of the Butajira district is almost representative of the climate of the nation: tropical lowlands, cooler highlands and temperate intermediate areas. Annual rainfall ranges between 900 and 1,400 mm, where the main rainy season is from June to September and “small rain” is common around March and April (Berhane, Wall et al. 1999). Malaria is endemic in the lowland areas.

Until the recent administrative restructuring split the Butajira district into two separate woredas (Meskan and Mareko), it was organized into four urban kebeles situated in Butajira town and 82 rural peasant associations. In the recent restructuring four of the rural peasant associations of the district have become part of another district, the Sodo district. Furthermore, some of the sub-districts are now included within the newly defined Silti zone. However, the name “Butajira district” will be used throughout this thesis for the sake of simplicity.

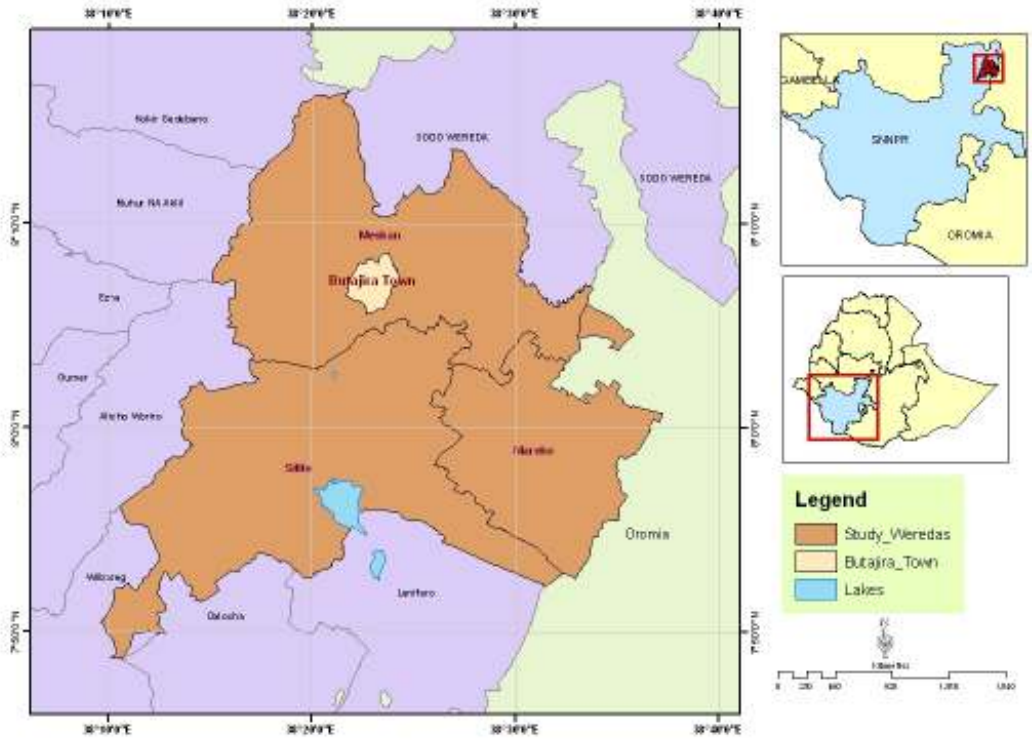

Figure 3.2: Location map of the Meskan Woreda, Mareko Woreda and Silti Zone

The people of Butajira district live on small-scale subsistence farming, cattle breeding and trading. Farmers grow maize, barley, wheat, sorghum, legumes and Enset (false banana). Enset is a drought-resistant plant whose farming is relatively easy but the process of converting it to edible food is labour intensive. For this reason women in the nearby villages work in groups whenever the processing of Enset becomes an agenda of a given household. Khat (Catha edulis) and chilli pepper are the main cash crops. Khat is a mild psychostimulant, about one quarter to one seventh the strength of amphetamine. It is widely grown and consumed in the area. Khat is used in religious rituals, particularly by followers of Islam.

Butajira district is one of the beneficiaries of Ethiopia's radical health service reform programme. At the peasant association level in rural settings and at a kebele level in urban settings there are now two female health extension workers (HEW) to serve a population of about 5000 persons (1000 house holds). HEW started their work in rural kebeles between 2004 and 2005 in rural kebeles and in 2010 in urban kebeles. The

primary responsibility of these HEWs is health promotion and prevention of illness through health education, house to house visits, facilitating immunisations and providing simple diagnostic and curative measures (for example for malaria). In each district there are a minimum of five health centres staffed by nurses and headed by health officers. There are two hospitals: a public hospital in Butajira town inaugurated in 2002 and a second hospital owned by a charitable organization (Project Mercy) located about ten kilometres south west of Butajira town. In all urban settings of the districts there are several low level private clinics which are staffed either by nurses or health assistants and provide non-surgical uncomplicated curative health services to the population. There is one psychiatric clinic, established by the ongoing research project of the Department of Psychiatry, Addis Ababa University. This clinic is staffed by two psychiatric nurses and it provides a service for patients coming from a vast area, from within and outside the Gurage and Silti Zones.

3.1.3 The Butajira Demographic and Surveillance Site (DSS)

In the absence of a national vital registration system, the sources of health information for the design and implementation of health policies in Ethiopia include but are not limited to, household surveys, disease surveillance and outbreak notification, censuses, data collected based on patient and service records, reporting from community health workers and health facilities, facility surveys, programme-specific monitoring and evaluation (e.g., TB, HIV/AIDS, EPI), and research and special studies. However, these sources have their own limitations including data quality, duplication and waste due to parallel health information systems, lack of timely reporting and feedback, centralization of information management without feedback to lower levels, variation in quality and completeness of reporting, timeliness of reporting and scope of coverage (e.g. may only cover government facilities).

With the above background information, the Butajira DSS was established in 1986 as an Ethio-Swedish collaborative research program by the departments of community health, AAU (now School of Public Health) and Epidemiology and Public Health Sciences, Department of Public Health, and Clinical Medicine, Umeå University, with the objectives of providing a base population, sampling frame, and infrastructure for

problem-oriented community based studies, in addition to providing epidemiological information that contributes to improved health management and decision-making (Shamebo, Sandstrom et al. 1993; Berhane, Wall et al. 1999). The DSS (figure 3.3**)** includes nine rural PAs from different ecological zones and one urban kebele in Butajira town. These study sites were originally selected using the probability proportional to size (PPS) method from the Butajira district (http://www.butajira.org).

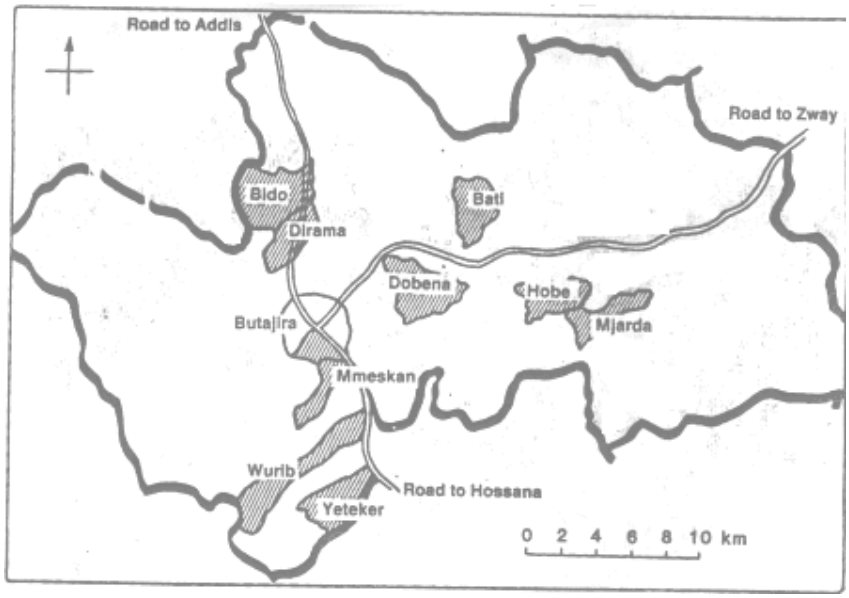

Figure 3.3: Location map of the 10 kebele/PAs included in the Butajira DSS

Starting with the base population determined after conducing the census in 1986, monthly registration of vital events had been the practice in the DSS between 1987 -1999 but was replaced by quarterly house-to-house visit in 2000 (Fantahun 2008). Each of the 10 DSS kebele/PAs has one or two full-time employed enumerators. These enumerators are residents of their respective kebele/PA who visit every household once every three months to collect data on vital events, allowing calculation of the exact population of the DSS. Health posts and elementary schools are located within each PA and are reasonably accessible to the residents (a maximum of 5km walk). During our data collection the majority of these health posts were staffed by one male community health agent (CHA) with three to six month training in primary health care. At present, however, every health post is staffed with two female health extension workers

3.2 The P-MaMiE study and growth monitoring of infants

3.2.1 Study design and participants

The Perinatal Maternal Mental Disorder in Ethiopia (P-MaMiE) study was established in the demographic and surveillance site (DSS) within Butajira in 2004 with the main objectives of investigating the sociocultural context, epidemiology and public health impact of perinatal common mental disorders (CMD) (Hanlon 2009). Two PAs of the DSS are from Mareko woreda, one PA is from Silti zone and the remaining six PAs and an urban kebele are from Meskan woreda. In 2007 the Butajira town had 33,393 inhabitants with a 1:1.03 male to female ratio, and the two woreda (Meskan and Mareko) together had 256,713 inhabitants (Central Statistical Agency 2008).

The P-MaMiE study is a population based prospective cohort study which has recruited its study participants from a predominantly rural population where undernutrition among children is a common phenomenon but the magnitude of perinatal CMD was unknown (Hanlon, Medhin et al. 2009). Eligible women were between the ages of 15 and 49 years, able to speak Amharic (the official language of Ethiopia), living in the DSS and in the third trimester of pregnancy during the study recruitment period (July 2005 to February 2006). The women were identified by DSS enumerators in the course of their 3-monthly surveillance interviews and, after giving informed consent, were interviewed by the P-MaMiE study female data collectors. Of the eligible pregnant women, 1065 (86.3%) were successfully recruited **(**figure 3.4**)**. Non-recruited women did not differ significantly from participating women in terms of age, religion, ethnicity, level of literacy, or location of residence.

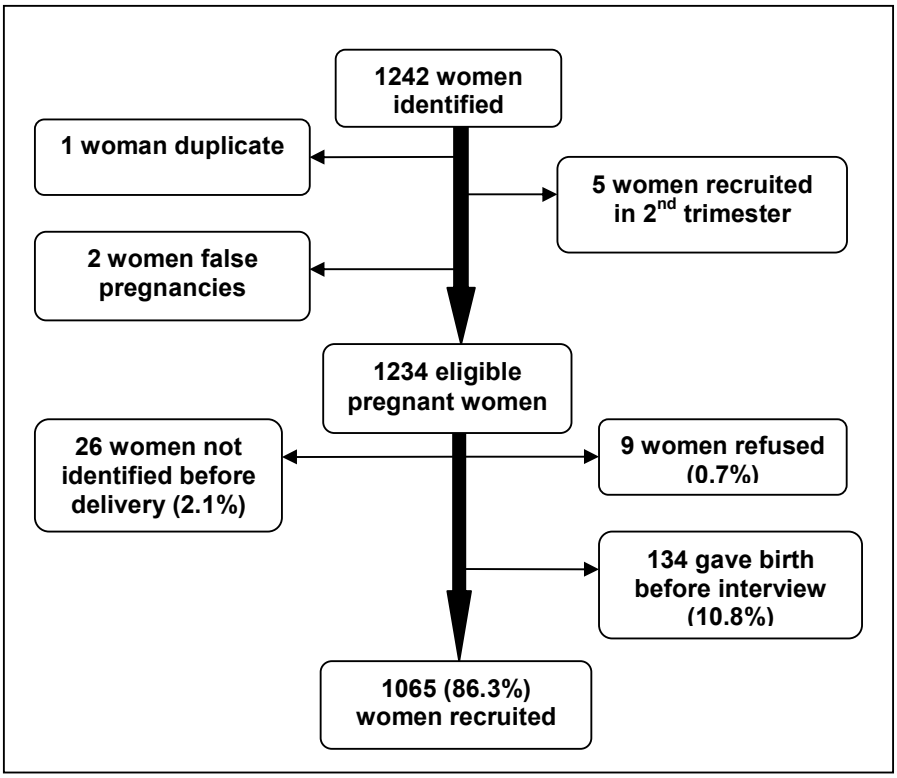

Figure 3.4: Flow chart showing recruitment of study participants into the P-MaMiE Study

After giving birth, mothers and their new babies were followed-up until the child was 18 months of age. During the whole follow-up period,mothers were assessed at four time-points (at recruitment, at birth, at two month postnatal and at one year postnatal) and infants' anthropometric measurements were monitored at five time points (at two, six, nine, 12 and 18 months of age).

3.2.2 Ethical considerations

Prior to the first interview, women were informed about the objective of the study and provided informed written consent. Arrangements were made for the study project to pay all health-related expenses of the mothers and children participating in the study from the day of recruitment to the first birthday of the infant. At recruitment women were given an identification card which allowed them to obtain healthcare on a credit basis (covered by

the project) and they were advised to deliver in health institutions or to consult the trained birth attendant in their respective PA. Traditional birth attendants were remunerated per each delivery they attended. In case the pregnancy ended in stillbirth or the infant died in the course of follow-up, the arrangement remained active for the mother. The study was granted ethical approval from the Ethics Committee of the Institute of Psychiatry, King's College London, and National Ethics Review Committee based in the Ethiopian Science and Technology Agency.

3.2.3 Data collectors

Ten female high school graduates, two with previous field data collection experience and one with primary school teaching experience, were employed during the preparatory period of the project to work fulltime on the P-MaMiE study. They were trained on the instruments to be implemented during data collection and on anthropometric measurement of mothers and infants. While they were in the third trimester of pregnancy participating women were interviewed in their house by these data collectors with very close day-to-day monitoring from the project coordinators. These data collectors continued to collect information through out the project follow-up period: (1) they interviewed the study participating women in the postnatal period at two and 12 month time points, (2) they measured growth of infants in one urban kebele in Butajira town where there was no Community Health Agent (CHA) available, (3) they worked as a back-up in collecting growth measurements in the remaining nine rural PAs whenever there was an unexpected delay from CHAs, e.g. due to mass vaccination campaign in the area, and (4) they did period checks of the quality of growth measurement of the CHAs.

One CHA and one trained traditional birth attendant (TBA) recommended by the woreda health bureau were selected from each PA and trained to accurately carry out anthropometric measurements of infants in their respective PAs. Using the CHAs had the advantage of capitalising on their previous experience of screening undernourished children and mothers on behalf of the woreda health bureao. There were no CHAs in two PAs and one CHA in the third PA was not willing to be involved in growth monitoring activities. Hence, two DSS enumerators and one high school graduate local farmer were involved in anthropometric measurements in these PAs. All these individuals were

allocated to attend every delivery of the study women in their respective PAs whenever possible or otherwise to visit the postnatal house at least within 48 hours of delivery. During their visit they recorded birth weight (if the visit was within 48 hours of delivery) and filled in a one page data collection form about the circumstances of delivery by interviewing the postnatal woman. If the postnatal house was left unvisited in the first 48 hours of delivery the one page data collection form was completed by the CHA/DSS enumerator at his/her next convenient time. If the delay happens to be more than three weeks, one of the P-MaMiE project data collectors collected information about delivery circumstances.

3.2.4 Data collection tools and methods

In this thesis we have used information (a) obtained from maternal reports through questionnaires administered by female data collectors, (b) about maternal anthropometric measurements collected by female data collectors, (c) about delivery circumstances obtained from maternal reports, the majority of them within 48 hours, and (d) about anthropometric measurements of infants measured by project data collectors, CHAs, TBAs or DSS enumerators.

Mental health measures

Reliability and feasibility of Comprehensive Psychopathological Rating Scale (CPRS) (Asberg and al 1978) was established as a rating of CMD in Ethiopia after training Ethiopian psychiatrists in use of the rating scale (Hanlon, Medhin et al. 2008). Using CPRS as a gold standard Self-Reporting Questionnaire (SRQ-20) was validated on perinatal women from the study area (Hanlon, Medhin et al. 2008). Following this CMD was measured during the third trimester of pregnancy and at two months postnatally using the locally validated Self-Reporting Questionnaire (SRQ-20) (Hanlon, Medhin et al. 2008). The SRQ-20 is composed of twenty yes/no items asking about the experience of depressive, anxiety, panic and somatic symptoms in the preceding 30 days (Beusenberg and Orley 1994). The SRQ-20 generates a continuously distributed scale score indicating overall psychological morbidity. In the current study area SRQ-20 showed acceptable convergent validity both as a linear scale and as ordered categories of

SRQ symptom burden: no symptoms (scored 0), low symptoms (one to five) and high symptoms (six and above) (Hanlon, Medhin et al. 2008) .

Risk factor questionnaire

To collect information during pregnancy on various risk factors we developed a questionnaire covering the following ten 10 domains: socio-demographics, household socioeconomic status, community and social support, time-framed life events, marital relationships, reproductive history, traditional attitudes regarding pregnancy, past or family history of emotional problems in the mother, substance use in the mother and her husband, and physical health of the mother. Questions for each domain were designed to cover the currently available literature and adapted from various sources including demographic and health survey (DHS) of Ethiopia(Central Statistical Authority (Ethiopia) and ORC Macro 2001). Some questions were also developed based on the qualitative study conducted during the preparatory phase of the P-MaMiE study (Hanlon 2009). A modified version of the same questionnaire was administered to the mother at two and at one year postnatal time points, with additional questions asking about the health of the infant and infant feeding practices.

Delivery circumstances

A separate questionnaire was administered to women at the time of measuring birth weight or in the first few days after birth in order to obtain information on stillbirths, prolonged labour, pre-lacteal feeding, whether colostrum was witheld, and the timing of initiation of breast-feeding. Timing of initiation of breastfeeding was recorded using three categories (< 1 hour, 1-8 hours, and > 8 hours following delivery). In the PAs where birth weight was not measured, the questionnaire was administered by the project data collectors, having been informed by the traditional birth attendant or DSS enumerators that the woman had given birth. As women in this setting were unable to report their gestation at delivery with any accuracy we were unable to assess preterm delivery. It was not feasible to use last menstrual period (LMP) to estimate gestational age in this cohort because the mothers could not reliably date their LMP. Ultrasound facilities were not available in the district and fundal height was unreliable due to the high proportion of multi-gravida women.

Birth weight

In six of the nine rural PAs, CHAs were trained to measure birth weight using SECA 725 scales measuring to an accuracy of 10g and kindly donated by UNICEF to the Meskan and Mareko Woreda Health Bureaus. The CHAs lived and worked in the PAs and were well-known to the women. After giving birth, participating women were requested to inform the CHA of the birth so as to enable the neonate to be weighed within 24 to 48 hours of birth. Birth weights were not measured in the remaining three PAs and in the urban kebele as no suitable health worker was available.

Growth monitoring of infants

As indicators of infant growth we used (1) weight measured using SECA 725 scales as above, and (2) length obtained using a locally adopted measuring board. These growth measurements were taken (1) by DSS enumerators in two PAs where there were no CHAs, (2) by high school graduate young farmer in one PA where the CHA and health extension worker were not willing to participate, (3) by CHAs in six PAs and (4) by project data collectors in an urban kebele. Using the CHAs had an advantage of their previous experience of screening undernourished children and mothers on behalf of the Woreda health bureau. Every person involved in growth monitoring was trained by the project co-coordinators (Girmay Medhin and Charlotte Hanlon) and one of the collaborators of the project to minimise inter-individual variability.

Quality control

Before the start of the actual data collection every effort was made in terms of instrument development, organizational setup of the project site and field office, screening of data collectors and training them in order to have a smooth and dedicated data collection process. Training was also conducted periodically over the course of the study period. The two lead project coordinators (Girmay Medhin and Charlotte Hanlon) established themselves on the project site 10 months before the start of the project. The two of them jointly administered the conduct of the whole study. During the field operation of the project, field activities were planned for a week and data collection was organized daily from the field office. Efforts were made not to overload data collectors so that the quality

of data collection would not be compromised. Quality of interview and measuring of growth were periodically monitored on site. Data collected by interviewing study participants was reported to the field office on a daily basis and anthropometric data was reported to the field office every Friday. The reason for the selection of Friday for the reporting of anthropometric data was to fit to the practice of DSS surveillance data reporting to the field office. The DSS enumerators submit their weekly surveillance data to the field office of the DSS located in Butajira town on every Friday.

Data was checked in the field by supervisors. Except for the first three months where the first and double data entry was managed by the two project coordinators, first data entry was managed by a data entry clerk and second data entry was managed by either of the two project coordinators using Epidata to minimize data entry error (Epidata (Version 3) 2003). First data entry was done within the first two days of its collection and double data entry was done within one week of the data collection. Incomplete or inconsistent data which was identified during data entry was corrected by sending the data collectors back to the homes of study participants

3.3. Variables used in this thesis

3.3.1 Outcome variables

In this thesis I have analysed six interrelated outcome variables, namely, length in centimetres, weight in kilograms, length-for-age z score, weight-for-age z score, stunting and underweight. In descriptive summaries we have also used weight-for-length z score and wasting. In traditional analysis, anthropometric measurements taken between five and seven months and between eleven and thirteen months were considered for as six and twelve month nutritional indicators, respectively. The standardize z-scores of infant length and infant weight measurements at each follow-up time were obtained using 2006 WHO reference population implemented in ANTRO 2005 (Monika, Elaine et al. 2006). Weight-for-length helps to identify acute undernutrition and length-for-age reflects reduced skeletal growth as the result of repeated undernutrition. Lower values of weight-for-length are termed as wasting (weight-for-length <-2) and that of length-for-age are called stunting (length-for-age <-2). Wasting is the result of short term measures from which children can

recover if fed and cared for appropriately. On the other hand, the prevalence of stunting is an indicator of accumulated undernutrition. Weight-for-age is a composite of the two other indices and its lower value (weight-for-age <-2) is termed underweight. A low weight-for-age (i.e. underweight) can be derived both from an insufficient length-for-age of the child or from a low weight-for-length of the child. Therefore, the index reflects both chronic undernutrition and acute undernutrition.

3.3.2 Main exposure variable

The main exposure variable was perinatal common mental disorders (CMD), used both as a continuous dimensional measure and as a categorical variable. When it was considered as a categorical variable the total score was dichotomised (SRQ-20 < 6 versus SRQ = 6), high scores indicating a high level of CMD. Three different exposure variables of CMD were considered:

- antenatal CMD prevalent cases,
- postnatal CMD - prevalent cases,
- a four level categorical exposure of CMD with the following categories - never had CMD (never exposed), antenatal CMD resolving after birth (antenatal only), incident postnatal CMD (postnatal only), and 'persistent' CMD (high SRQ-20 score antenatally and postnatally)

3.3.3 Other risk factors or confounding variables

Composite variables (or scales)

Including numerous individual items in a model enables measurement of the effect of the specific detailed item on the outcome which might be helpful in designing interventions. On the negative side, there is a higher likelihood of getting false positive associations as the result of multiple tests. To overcome the latter problem without significant compromise of the first issue, the following four composite scores were created from items defining the same concept by adding responses of identified variables giving equal weights to each item. Each composite variable was used as a continuous variable:

(1) Poor sanitary conditions: not having a toilet facility, not having safe water, disposing of rubbish on the field). These three variables were aggregated since all of them are known risk factors for undernutrition in Ethiopia, although their internal consistency measure was relatively low (Cronbach alpha = 0.49)

(2) Maternal empowerment: decision making power reflecting whether the woman is allowed to act without asking permission from her husband (to sell crops, to spend household money, attend meetings like women's groups, buy medication for herself or her children, to go to a health institution for health education or for medical examination). The scale has Cronbach alpha value of 0.93.

(3) Availability of support to the mother: able to visit friends, enough help at home, enough help in looking after children other than index child, enough help from husband and no violence towards her. This scale has Cronbach alpha value of 0.47 which is relatively low but these items measure quite different sources of support and we would not expect them to correlate highly.

(4) Poverty index: non-literate wife, non-literate husband, do not own radio, do not own bed, do not possess valuable goods like gold and jewellery, possess house, have large animals, have small animals, animals spending the night within the living house, house they live in has a window. We formed this scale by an exploratory factor analysis starting with more than 30 characteristic variables and tested the resulting three factors with confirmatory factor analysis. The factors and their indicators were then modified to attain convergence and also meaningful factor loadings. We then defined the index as the sum of individual items with equal weight after obtaining a one factor model with meaningful factor loadings in terms of expected sign, statistical significance and width of confidence intervals. At this stage we preferred to give equal weights for the identified variables to create a composite score since we could not find convincing evidence to prioritize one over the other in terms of its potential effect on infant nutritional status in this setting. The scale has Cronbach alpha value of 0.73 which shows an acceptable level of internal consistency.

Other potential risk factors of infant growth

The following characteristics were treated individually within logistic and linear regression analysis while modelling nutritional status of infants: residence (urban or rural), number of other children under the age of five years(0, 1, 2+), age of father, age of mother, height of mother, mid-upper arm circumference of mother, marital status (polygamous versus non-polygamous), substance use as a binary (using of khat (*cata edulis Forsk*) at least weekly or drinking alcohol at least weekly), at least one obstetric complication (prolonged labour (>24 hours), assisted delivery (normal vaginal delivery versus instrumental/ Caesarian section), post-partum haemorrhage, post-partum fever), infant gender, immunisation status of infant at two months (yes/no) as an indicator of maternal health seeking behaviour, history of infant's illness in the first two months of life to the extent that the mother thought the baby was going to die (yes/no), birth weight (low (BW < 2500g) , normal (BW >=2500g), not measured) and early infant feeding practices. Particular focus was given to characteristics that described how the child was fed in the first two months of life. Specific interest was whether the infant was exclusively breast fed in the first two month as reported by the mother at two month follow-up (yes/no), if pre-lacteal food was given (yes/no), if colostrum was given (yes/no) and the timing of initiation of breast feeding (less or equal to one hour versus more than one hour).

3.3 Data analysis

3.3.1 An overview of steps followed during data analysis

Data analysis in this thesis included singleton infants who had growth measurements from at least one time-point and their mothers. The analysis involved a descriptive summary of various variables, bivariate associations, cross-sectional modelling of growth measurements of infants taken when infants were 2, 6 and 12 months of age using logistic and linear regression depending on the scale of each growth outcome under investigation, and longitudinal modelling of infant growth outcome measures recorded at up to five time points using multilevel and latent growth modelling techniques. The majority of potential risk factors for poorer infant growth were extracted from maternal

responses collected at recruitment when mothers were in the third trimester of pregnancy. Additional risk factors were extracted from the information collected shortly after delivery or at two months postnatal.

Growth outcome data was used from different time points depending on the type of modelling. When logistic regression and linear regression were used as the modelling techniques we used infant growth data from when infants were two, six and 12 months of age. Whenever the timing of the growth measurement for two and six month time-points deviated by more than the targeted two weeks window either side, we expanded the age window so that any measurement before three months of age were considered as "two months" and any growth measurement between five and seven months of age were categorized as "six months". For the latent growth modelling we created an additional four month time-point by grouping together all infants that had their growth measurement taken between three months of age and five months of age. Since multilevel growth modelling uses the exact time of growth measurement as the dependent variable there was no need to regroup growth measurements for this analysis.

In the following sections we will describe the procedures that we have followed to analyze the data. The results are summarized in chapters 4-7.

- **Traditional modelling techniques**: Under this subheading we will briefly describe the type of descriptive methods used, variables used in logistic and linear regression and how these variables were coded.
- **Multilevel growth modelling:** Under this subheading we will briefly introduce the type of polynomial growth models that were used to select the best fitting unconditional growth models for each outcome and the type of conditional multilevel growth models fitted to the infant growth data.
- **Latent growth modelling:** Under this subheading unconditional and conditional latent growth modelling will be introduced, linked with the type of hypothesised growth models to be estimated using infant growth data. A description of unconditional growth models will be presented for continuous and binary outcomes. A more detailed account will be given of unconditional latent growth models applied to binary outcomes as this technique is less familiar compared to modelling of continuous outcomes. A description of conditional latent growth

models will be linked to models that are hypothesised to evaluate unadjusted, partially adjusted, mediated and fully adjusted effects of maternal CMD on infant growth.

3.3.2 Traditional modelling techniques

Before embarking on logistic regression and linear regression to evaluate the effect of CMD and other factors on infant growth, the data were cleaned and descriptive analysis carried out: simple frequencies, cross tabulation of categorical variables, descriptive summaries and graphical presentations of continuous growth outcomes and assessment of the assumptions required for the application of parametric tests. This was then followed by bivariate analyses: chi-square to examine associations between categorical variables, t-tests for comparison of means, and unadjusted estimates of effects (regression coefficients or odds ratio) to describe the association between a given growth outcome and different exposures. To maximise the use of available data, birth weight (BW) was included in all regression models as a three category variable (normal (BW $\geq$ 2500g), low birth weight (BW <2500g) and birth weight not recorded)

While summarizing the data means and proportions were used to describe continuous and categorical characteristics, respectively. Independent sample t-tests were used to compare the mean score of growth indicators of infants born to mothers with and without a high level of CMD. The proportions of undernourished infants among those born to mothers with and without high levels of CMD were compared using Fisher's exact test. The independent effect of CMD on infant growth was evaluated by defining three main exposure variables: (a) antenatal prevalent case, (b) postnatal prevalent cases, and (c) a four level categorical exposure variable ("no exposure at both time points" (reference), only antenatal exposure, incident postnatal, and "chronic" or persistent exposure) of CMD. Taking each of the three CMD exposures in turn, the association with infant growth was investigated with logistic regression for binary outcomes (undernourished versus well-nourished) and linear regression for continuous growth outcomes (standardized and unstandardized growth measures).

In the process of modeling each growth outcome at each time point (two , six and 12 months) to evaluate the effect of CMD three steps were followed:

- bivariate regression taking one of the three CMD exposure variables at a time,
- multivariable regression adjusting for the effect of CMD on an outcome for a given domain of covariates (household characteristics, child characteristics, maternal characteristics, or infant feeding practices), and
- multivariable regression fully adjusting the effect of CMD for all covariates.

Unadjusted and adjusted odds ratios from logistic regression and unstandardized regression coefficients from linear regression, with corresponding 95% confidence intervals, were used to assess statistical significance and the magnitude of effects.

Similarly, in evaluating potential risk factors of infant growth other than CMD at two, six and 12 months of infant age, logistic regression for binary outcomes and linear regression for continuous outcomes were used. Unadjusted and adjusted odds ratios from logistic regression and unstandardized regression coefficients from linear regression with corresponding 95% confidence intervals were used to assess the significance and the magnitude of the effect of a given exposure. In building the fully adjusted model, the following two steps were followed: (1) estimate the fully adjusted model for all variables hypothesized as risk factors except for the four variables which represented infant feeding practices, (2) to obtain the independent effect of each "feeding practice", we re-estimated the model in (1) including also the feeding practice variable of interest. This generated (a) the effect of each variable adjusted for all others except feeding practice variables, (b) the effect of each feeding practice variable adjusted for all other variables but not for the remaining feeding practice variables. We did not adjust each infant feeding practice for the others because of collinearity and to enable us to obtain the independent effect of each feeding practice on infant growth. All data analysis was carried out using STATA version 10 (StataCorp 2007) with the probability of type 1 error set at 5%. .

3.3.3 Multilevel Growth Modelling

3.3.3.1. Overview of multilevel growth models

This technique involves fitting two groups of models for each growth outcome, namely unconditional and conditional multilevel growth models. Within each group of these models, hierarchical models were fitted and then compared to assess the degree to which the overall model fit was improved as well as to test the statistical significance of each regression coefficient (i.e. fixed effects).

(a) Unconditional multilevel growth model

An unconditional multilevel growth model was used to select the best fitting baseline (i.e. within infant over time) growth model for each of the six growth outcomes. This class of growth model has the following general form:

$$\begin{cases} E(y \mid x) = \mu \\ g(\mu) = \sum_{k=0}^{K} \beta_k t^k + u_{0i} + u_{1i} + \varepsilon_{ij} \end{cases}$$

where

- E(y|x) stands for the expectation or mean of a given growth outcome y conditional to the values of x
- $g(\mu)$ stands for the general link function (eg. identity link for normally distributed growth outcomes and logit link for binary growth outcomes),
- $\beta_k, k = 0,1,..K$ are regression coefficients,
- K is the highest degree of polynomial to be used in modelling the growth function of an individual infant over time,
- t represents some form of time (eg. exact time or centred value) at which the growth measurement was recorded,
- u_{0i} is the random intercept and u_{1i} is the random slope (they are measures of discrepancies between the true overall parameter values of all infants – intercept and slope – and within individual infant growth parameters of i^{th} infant – intercept and slope). These two random parameters are assumed to have a bivariate normal distribution.

- ε_{ij} is the within i[th] infant residual at j[th] measurement time point, and assumed to be independent of (1) the two random coefficients (i.e. u_{0i} and u_{1i}), and (2) the timing at which infant growth was measured.

We started modelling of each growth outcome of infants by fitting the simplest model that can be obtained if we let K = 0:

$$\begin{cases} E(y/x) = \mu \\ g(\mu) = \beta_0 + u_{0i} + \varepsilon_{ij} \end{cases}.$$

This growth model assumes that there are no changes in the outcome of interest over time (i.e. the overall mean of infants is β_0 across time) but individual infant can have different mean value of the growth measurement under investigation (eg. length) which is equal to $\beta_i = \beta_0 + u_{0i}$. The model was then compared with its counterpart that assumes the same mean value for all infants. For example the comparison model for length of an infant was the one that could be obtained by fitting an ordinary regression model without a slope (i.e. length = $\beta_0 + \varepsilon_{ij}$).

Provided the random intercept model was superior to the one which does not allow the means to vary across infants, the next logical step was to relax the model assumptions by allowing the linear slope to be in the model. This involves letting K = 1 which gives the following two unconditional multilevel growth models:

- $E(y \mid x) = \mu$ and $g(\mu) = \beta_0 + \beta_1 t + u_{0i} + \varepsilon_{ij}$ which allows all infants to have the same slope (β_1 is the same for all infants) but varying intercept (i.e. $\beta_{0i} = \beta_0 + u_{0i}$).
- $E(y \mid x) = \mu$ and $g(\mu) = \beta_0 + \beta_1 t + u_{0i} + u_{1i} + \varepsilon_{ij}$ which allows both intercept ($\beta_{0i} = \beta_0 + u_{0i}$) and slope ($\beta_{1i} = \beta_1 + u_{1i}$) to vary across infants. In the equation of $g(\mu)$ the random slope indicator u_{1i} is a function of time (i.e. $u_{1i} = \xi_{1i} * t$) allowing the slope to vary with time

To proceed to the quadratic unconditional growth model which does not allow the random curvature, the value of K = 2 and its general form is as follows:

$E(y \mid x) = \mu$ and $g(\mu) = \beta_0 + \beta_1 t + \beta_2 t^2 + u_{0i} + u_{1i} + \varepsilon_{ij}$ where all the parameters are as defined before. A new parameter in this model is β_2 which represents the non-random slope of the quadratic term of time or a measure of curvature.

Theoretically, the quadratic growth function can also have related a random component (it would be u_{2i} in the equation). However, our data has a maximum of five measurements per person and it becomes difficult to estimate such a large number of parameters from a limited number of data points. That means we should limit the number of parameters to be estimated from the data.

In summary, therefore, we fitted the above unconditional growth models in the following hierarchical order:

(a) random intercept but no slope in the model,
(b) random intercept and fixed slope,
(c) random intercept and random slope, and
(d) random intercept and random slope but fixed rate of curvature (quadratic term).

Improvement in model fit as the result of fitting a less restricted model, which is a one step higher in the hierarchy, compared to a restricted model was assessed using change in deviance and the Information Criteria (AIC and BIC). Statistical significance of fixed effects (i.e. the $\beta_i, i = 0,1,..K$) was assessed using the Wald statistic.

Following the fit of each unconditional growth model for a given growth outcome the overall variability in that outcome variable was partitioned into its various components (estimates of random coefficients and within infant random error) and their corresponding variances and covariance were estimated. This partitioning was used to check if there was significant variability in the random coefficients that need to be explained by infant level variables (i.e. parental, household, and infant characteristics) implying the need for conditional multilevel growth models. In the case of the random intercept growth model where the outcome was continuous it was also possible to estimate intra-class correlation by taking the ratio of the variance components that correspond to the random intercept and the overall variance (i.e. $\rho = \frac{\text{var}(u_{01})}{\text{var}(u_{01}) + \text{var}(\varepsilon_{ij})}$).

However, in the random coefficients model total variance varies as a function of time which makes it difficult to define intra-class correlation.

(b) Conditional Multilevel Growth Models.

By conditional multilevel growth modelling we mean a model with predictors other than time in the best fitting individual growth model. Additional predictors are included in an individual level growth model to explain some of the variability observed in the random components (random intercept or random slope) of the unconditional growth model. For example, let us assume a normally distributed continuous growth outcome, say y, whose best fitting unconditional within-infant growth model is linear with random intercept and random slope. If we consider just one predictor variable, say x, of the random coefficients (random intercept and random slope) then the conditional multilevel growth model will have the following three equations:

$$y_{it} = \beta_{0i} + \beta_{1i}t + \varepsilon_{ij}$$
$$\beta_{0i} = \beta_0 + \gamma_0 x + u_{0i}$$
$$\beta_{1i} = \beta_1 + \gamma_1 x + u_{1i}$$

The first equation ($y_{it} = \beta_{0i} + \beta_{1i}t + \varepsilon_{ij}$) represents the within individual growth model and the second two equations are regressing infant specific intercepts (β_{0i}) and infant specific slopes (β_{1i}) collected from individual infant's linear growth functions on the predictor variable (i.e. x in this cases). The intercepts of the second two equations, β_0 and β_1, are interpreted as the initial value and rate of change of an average infant, respectively. Similarly, γ_0 is interpreted as the effect of x on the initial growth of an average infant and γ_1 is interpreted as the effect of x on the rate of growth of an average infant.

Examining the above three sets of equations and replacing β_{0i} and β_{1i} in the first equation by their predicting equations presented in the second and third equations gives the following composite equation:

$$y_{it} = (\beta_0 + \beta_1 t) + (\gamma_0 x + \gamma_1 x * t) + u_{0i} + u_{1i}(= \xi_{1i} * t) + \varepsilon_{ij}$$

where β_0 and β_1 keep their previously described meanings. For example in modelling our infant growth outcome data, if a significant effect of a given covariate (e.g. perinatal maternal CMD) on the rate of infant growth is hypothesized its interaction term with time should be included as one of the covariates at the stage of model fitting.

3.3.3.2. Application of multilevel growth modelling to infant growth data

In modelling infant growth data using the multilevel growth modelling technique the best fitting unconditional polynomial growth model was identified for each growth outcome variable and predictors of the random components were introduced at different steps. This stage of modelling involved five steps depending on which covariates we were focusing on:

- unadjusted effect of CMD on growth parameters,
- unadjusted effect of covariates other than CMD on the initial value (intercept) and on the rate of growth (slope),
- adjusted effects of all covariates other than CMD on initial value and adjusted effects of selected covariates on the rate of growth over time,
- adjusted effect of CMD on initial value of growth.

The reason for treating CMD differently from the rest of covariates was to properly address the main objectives of the study

The conditional multilevel growth model that we have used in this study has the following general form:

$$E(y/x) = g(\mu)$$

$$\mu = \sum_{p=0}^{P} \gamma_p Z_p + \sum_{k=0}^{K} \beta_k t^k + \sum_{v}^{V} \sum_{k=0}^{K} \lambda_v Z_v t^k + u_{0i} + u_{1i} + \varepsilon_{ij}$$

where the first summation corresponds to newly introduced predictors and their corresponding regression coefficient, the second summation corresponds to the different parts of the polynomial function identified in the unconditional multilevel growth model above and their corresponding regression coefficients, the third set of summations corresponds to the interaction of newly introduced covariates with time and their corresponding regression coefficients implying the effect of covariates on the rate of growth (i.e. slope), and the last three terms correspond to the overall residual of the model representing random coefficients (the first two components) and individual level residual (last component).

3.3.4 Latent Growth Modelling

3.3.4.1 Overview of LGM

Use of LGM to evaluate the effect of maternal CMD on infant anthropometric growth involved fitting two sets of models, namely, unconditional and conditional LGMs (Bollen and Curran 2006). Unconditional LGM was used to decide on the best fitting functional form of growth outcome over time and conditional LGM was used to investigate the effect of maternal CMD and other covariates on growth parameters of the best fitting unconditional LGM. The functional form of unconditional LGM depends on how the timing of growth measurement is parameterized while the interrelationship of variables in the conditional LGM depends on the research question under investigation. The scale of observed growth outcome variable, its distribution, and the assumption made about its underlying latent variable are some of the factors that affect model specification, estimation and interpretation of model parameters. Like any other SEM techniques (Bollen 1989) LGM imposes the assumption of a normal distribution on a continuous measured outcome variables (Bollen and Curran 2006). If the outcome variable is categorical, implying violation of the normal distribution assumption (e.g. stunting of infants), inconsistent estimates of mean vector and covariance matrix application of LGM requires special consideration (Mehta, Neale et al. 2004) as described in the background chapter of this thesis.

In the following sections I will describe the procedures that were followed in order to fit LGM to infant growth data and the details of models estimated in the following order

(a) Unconditional LGM for continuous growth outcomes ,
(b) Unconditional LGM for binary growth outcomes,
(c) Unadjusted effect of CMD on infant growth,
(d) Partially adjusted effect of CMD on infant growth
(e) Mediated effects of CMD on infant growth
(f) Fully adjusted effect of CMD and other factors on infant growth

3.3.4.2 Unconditional LGM

a. Unconditional LGM for continuous growth outcomes

For a continuous infant growth outcome, say Y, measured at six follow-up time points the unconditional LGM can be expressed in matrix form (Bollen and Curran 2006) as

$$Y = \Lambda\eta + \varepsilon$$

where

- Y is a 6 x 1 vector of repeated growth measures,
- Λ is a 6 x m matrix of factor loadings,
- η is an m x 1 vector of latent factors representing parameters underlying the growth trajectory,
- m is the number of latent factors hypothesized to characterize the underlying unconditional latent growth model
- and ε is a 6 x 1 vector of residuals

Linear unconditional LGM is characterized by two latent factors (i.e. initial value and slope) and with the assumption of six repeated growth measurements per infant the elements of the matrix $Y = \Lambda\eta + \varepsilon$ are

$$\begin{pmatrix} y_{i1} \\ y_{i2} \\ y_{i3} \\ y_{i4} \\ y_{i5} \\ y_{i6} \end{pmatrix} = \begin{pmatrix} 1 & 0 \\ 1 & 1 \\ 1 & 2 \\ 1 & 3 \\ 1 & 4 \\ 1 & 5 \end{pmatrix} * \begin{pmatrix} \alpha_i \\ \beta_i \end{pmatrix} + \begin{pmatrix} \varepsilon_{i1} \\ \varepsilon_{i2} \\ \varepsilon_{i3} \\ \varepsilon_{i4} \\ \varepsilon_{i5} \\ \varepsilon_{i6} \end{pmatrix}$$

This implies that each observed growth measure of an infant at a given follow-up time (y_{it}, i = 1, …n and t = 1, 2, ..6) can be described as the weighted combination of a random intercept α_i and a random linear slope term β_i plus an infant and time specific residual ε_{it} .

In general a two factor unconditional LGM model is described as $y_{it} = \alpha_i + \lambda_t\beta_i + \varepsilon_{it}$ where α_i and β_i stand for two latent variables which are commonly known as intercept

and slope, respectively, λ_t ($t = 1, 2....T$) are values linked to the time at which the growth outcome were measured, and t is the indicator of time at which infants' growth was recorded. It is a common practice to centre time in order to give a meaning for α_i. In this thesis, for example, I subtracted two from the time of growth assessment (i.e. time was centred at two months) to assign the meaning of "starting value" to the two month time point so that the intercept term (i.e. the expected value of α_i) in the regression equations will be interpreted in the same way. The two latent variables have their corresponding residual terms which are allowed to be correlated (i.e. $\text{cov}(\alpha_i, \beta_i) = \sigma_{\alpha\beta}$). Depending on how the time of growth outcome measurement is parameterized, the $\lambda_t's$ will take constant values or one estimated from the data simultaneously with other parameters of the model.

Three unconditional LGMs (Bollen and Curran 2006) described in the following pages mathematically in models 1-3 and graphically in figures 3.5 – 3.7, were fitted to the repeated continuous measurement of infant growth (i.e. length, weight, length-for-age, weight-for-age) measured in the rural villages of Butajira. The growth measurements used in this modelling were recorded when the infants were two, four, six, nine, 12 and 18 months of age. The three candidate growth models were

- linear LGM characterized by an intercept term and a linear slope term,
- quadratic LGM characterized by intercept, linear and quadratic terms, and
- non-linear LGM which is not quadratic characterized by intercept term, slope term, and four freely estimated loadings which allow the non-linear growth structure of LGMs over time.

Each of the above models has two groups of equations defining within-individual infant growth over time (level 1) and growth of infants at a group level (level 2).

Taking infant length as an example of a continuous growth outcome measure, level 1 and level 2 growth equations for the three LGMs are described below:

Model 1: (Linear)

$$length_{it} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it} \quad \text{……..} \qquad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \quad \beta_i = \mu_\beta + \zeta_{\beta_i} \quad \text{…….} \qquad \text{Level 2}$$

where

- μ_α is mean growth of an average infant at initial time,
- μ_β is mean rate of growth of an average infant per unit change in time,
- ζ_α and ζ_β are random errors of initial value and rate of change, respectively, which are assumed to have a bivariate normal distribution.

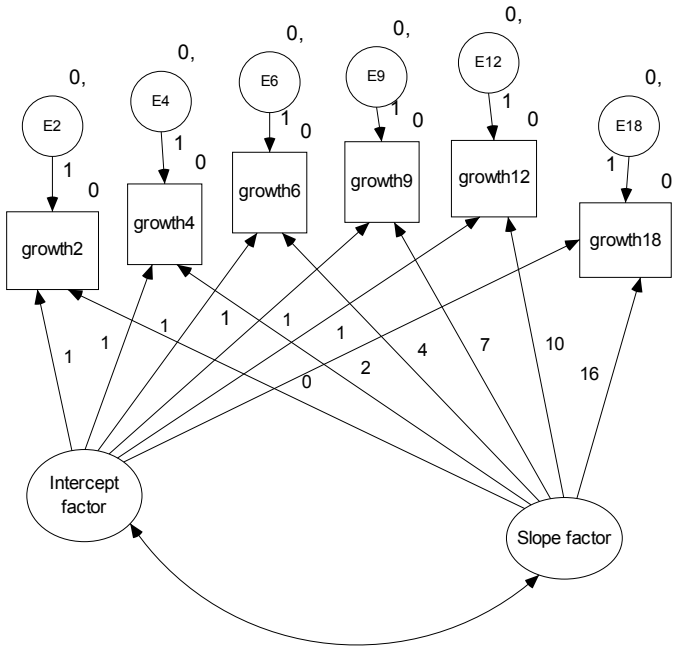

Figure 3.5: Hypothesized unconditional linear LGM for infant growth in Butajira, Ethiopia

Model 2: (Quadratic)

$$length_{it} = \alpha_i + \lambda_t \beta_{1i} + \lambda_t^2 \beta_2 + \varepsilon_{it} \ \ldots\ldots\ldots\ldots \quad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \ \ \beta_{1i} = \mu_{\beta_1} + \zeta_{\beta_{1i}} \ and \ \ \beta_2 = \mu_{\beta_2} \ldots\ldots \quad \text{Level 2}$$

where

- μ_α is the mean growth of an average infant at the initial time-point,
- μ_{β_1} is the mean rate of growth of an average infant at the initial time-point,
- μ_{β_2} is the mean rate at which the growth rate of an average infant (i.e. μ_{β_1}) changes per unit change in time (it remains the same for all infants),
- ζ_{α_i}, and $\zeta_{\beta_{1i}}$ are random errors of the initial value and rate of change, respectively, which are assumed to have bivariate normal distributions.

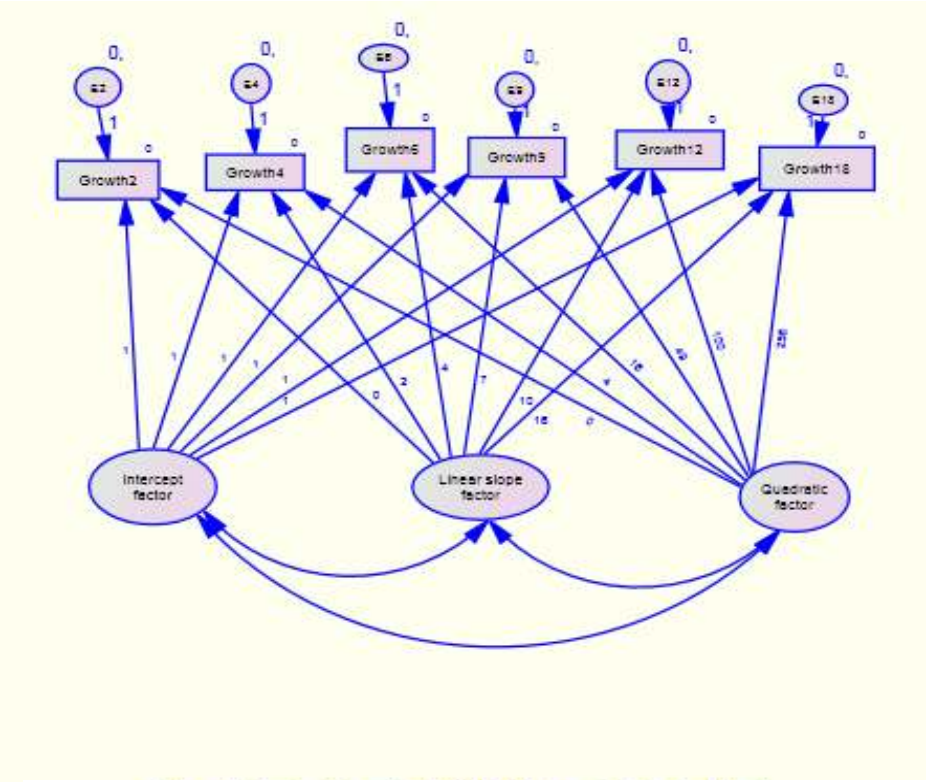

Figure 3.6: Hypothesized unconditional quadratic LGM for infant growth in Butajira, Ethiopia

Model 3: (Non-linear)

$$length_{it} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it} \quad \ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots \quad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \quad \beta_i = \mu_\beta + \zeta_{\beta_i} \quad \ldots\ldots\ldots \quad \text{Level 2}$$

where

- μ_α is the mean growth of an average infant at the initial time-point,
- μ_β is the total mean change in growth of an average infant between the initial time-point and the next time-point, fixed to give a unit of scale for time,
- λ_t, t=1, 2, ..T are loading of observed growth measures to the mean change factor (two of them have fixed values of 0 and 1 and the rest are to be freely estimated from the data)
- ζ_{α_i} and ζ_{β_i} are random errors of the initial value and total mean change per unit of time, respectively, which are assumed to have bivariate normal distributions.

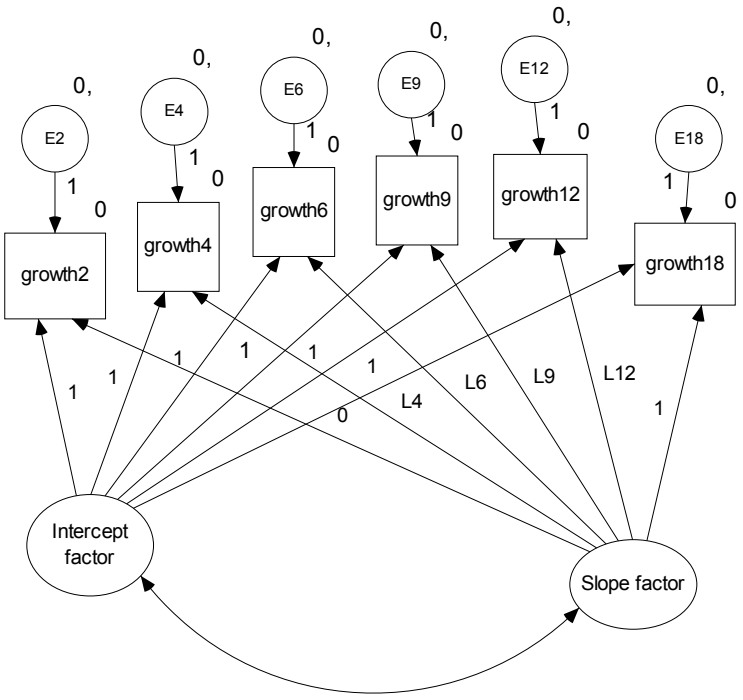

Figure 3.7: Hypothesized unconditional non-linear LGM for infant growth in Butajira, Ethiopia

In these models the values of λ_t determines how time of measurement or passage of time should be coded. In models 1 and 2 the values of λ_t have values equal to (age-2) where "age" corresponds to the time of each growth measurement and in model 3 only two of the λ_t's are fixed to 0 and 1, and all the rest λ_t's are estimated freely from the data simultaneously with other parameters of the model. This way of coding time facilitates the shape of growth within an individual to be linear (i.e. Model 1), quadratic (Model 2) or non-linear (Model 3) over time. Subtracting two from the time of growth assessment assigns two months to the starting value for growth and makes the intercept term meaningful (i.e. the expected value of α_i) in the regression equations. Although any two time-points can be used as the reference interval for the change in Model 3 (Bollen and Curran 2006) we took two months as the starting time point so that $\lambda_1 = 0$, the eighteen month assessment as the point of final assessment ($\lambda_6 = 1$), and we left all the remaining measurement occasions (i.e. $\lambda_2, \lambda_3, \lambda_4, \lambda_5$) to be estimated as free model parameters. The way we parameterized Model 3 enable us to interpret any change between the starting

point (i.e. the two month time point in our case) and the time under investigation as the proportion of change over the whole follow-up period.

As it is described in figures 3.5-3.7 the loadings of the intercept terms from all growth measurement time points are fixed to one in each of the three unconditional LGMs to facilitate model identification. In linear and quadratic LGMs the loadings of the linear slope term that comes from the first growth measurement time (i.e. age -2) are fixed at zero to give the interpretation of mean initial value for the intercept term and all other loadings are fixed at (age -2) to allow linear change over the age range under investigation. In the quadratic LGM the loadings of the quadratic term are fixed at $(age-2)^2$ to allow estimation of a quadratic slope term from the data. In the non-linear LGM the loadings of the slope term at two months of age (initial time) and eighteen months (final time) are fixed at 0 and 1, respectively, to give a scale for time and the remaining loadings are left as free model parameters to be estimated from the data. The intercept of all observed growth measures are fixed at zero to allow estimation of mean values for the latent factors characterizing each of the LGMs.

The three alternative unconditional LGMs were fitted to the four continuous growth outcome variables using AMOS 7 (Arbuckle 1995-2006) and data were used from six time points representing age at assessment (i.e. two, four, six, nine, 12 and 18 months). The modelling started with the simplest model and moved step by step to the more complex model. We started by fitting an unconditional LGM with linear and non-linear options for the slope term knowing that specification of linear option is nested within the non-linear option. The hypothesis was that specification of a non-linear slope would significantly improve the overall model fit over the assumption of linear growth. Covariance terms were included between slope and intercept terms. In each of these unconditional models, significant serial correlation among measurement errors was hypothesized and their statistical significance tested. Non-significant covariance among measurement errors was gradually dropped until only significant paths remained, as long as their exclusion did not significantly affect measures of overall fit indices. The choice for best fitting model among linear, quadratic and non-linear was based on models that assume independence of residuals. Further improvement of model fit was assessed by comparing the two models that differed on the independence of residuals assumption.

The intercepts represent average level of measured growth outcome at two months of age (eg. average length of infant at the age of two months). The slope term in unconditional LGM represents the change in target growth outcome for an average infant (a) per unit time which remains the same over the study period in case of linear growth model, (b) per unit time at two months whose magnitude is modified by quadratic term as time increases in case of quadratic growth model, and (c) over the whole study period (i.e. from two to 18 months) for the non-linear growth model.

b. Unconditional LGM for binary growth outcomes

Like any other structural equation modelling techniques, LGM assumes that the outcome of interest is continuous and normally distributed (Bollen and Curran 2006). In case of binary outcome variables (e.g. stunting and underweight of infant in Butajira) the variables are obviously not continuous and a higher degree of kurtosis is very likely, the identity link is not appropriate in modelling within-individual growth and the model implied variance-covariance matrix obtained using the identify link for a within-individual growth model is not a consistent estimate of the observed variance-covariance matrix of the binary outcome variables (Mehta, Neale et al. 2004; Bollen and Curran 2006). This problem was addressed by assuming underlying normally distributed latent variables for each binary outcome and (1) by defining a threshold that linked observed binary outcome with the underlying latent variable, (2) estimating tetrachoric variance-covariance matrix from the observed binary variables, and (3) using estimating methods which result in consistent estimates of the variance-covariance matrix (Olsson 1979; Muthen and Muthen 1998-2009; Mehta, Neale et al. 2004).

For the two binary growth outcomes (i.e. stunting and underweight of infants) within-subject growth models were specified, similar to continuous growth outcomes. Similarly, the same steps were followed in model fitting, goodness-of-fit evolution and selection of the best fitting-within individual growth model. The difference was that a probit link was used for binary outcomes in place of the identity link which is more appropriate for continuous outcomes (Muthen and Muthen 1998-2009; Muthen and Asparouhov 2002). Nutritional indicators recorded at two months of infant age were considered as the initial value, and attained by centreing time at two months. These models were estimated using Mplus software(Muthen and Muthen 1998-2009) while the models for continuous

outcome variables were estimated using AMOS software (Arbuckle 1995-2006). The default settings of the MPlus software were used. These allow estimation of the mean of the slope term as a free parameter, fix the mean of the intercept term at zero and allow estimation across a time-invariant threshold.

The estimate of the threshold obtained from Mplus software (Muthen and Muthen 1998-2009) is equal to the negative of the mean intercept term and it is used to calculate the probability of "success" (i.e. probability of stunting or underweight) at the initial time point (i.e. at two months in this thesis) from the cumulative distribution of the standard normal distribution. In other words, the mean value of the intercept term is the probit of the model predicted prevalence of undernutrition at two months of age. Being a multiplier of time, the freely estimated mean slope is interpreted as the overall change in the probit of stunting from two to 18 months. It is also a multiplier of the freely estimated loadings at four, six, nine and 12 months and useful in interpreting the overall change in the probit of stunting between two and any one of these time points.

3.3.4.3 Conditional LGM: modelling the effect of CMD on infant growth

3.3.4.3.1 Summary

Antenatal and postnatal CMD are the two main exposure variables in conditional LGMs that were used to evaluate the effect of perinatal CMD on infant growth. The objectives were addressed by investigating the effect of CMD on parameters of LGM underlying each anthropometric growth outcome, namely, length, weight, length-for-age, weight-for-age, stunting and underweight. After selecting the best fitting unconditional LGM for each growth outcome, the effect of CMD was investigated in partially adjusted, mediated and fully adjusted conditional LGMs which involved the following three steps:

(1) estimating the effect of antenatal and postnatal CMD on the parameters of the LGM underlying each growth outcome after adjusting for each other and allowing postnatal CMD to have a mediating role between antenatal CMD and infant growth,

(2) estimating the effect of antenatal and postnatal CMD on the parameters of LGM underlying each growth outcome after taking account of the potential mediating role of birth weight, early infant feeding practices and early infant illnesses, and

(3) estimating the effect of antenatal and postnatal CMD as described in (2) after adjusting for selected baseline characteristics of the mother and the household, early infant feeding practices of the mother and selected characteristics of the index infant.

3.3.4.3.2 Partially adjusted effects of CMD on parameters of unconditional LGM

In a partially adjusted conditional LGM, the only predictors of parameters of unconditional LGMs are antenatal and postnatal CMD (figures 3.8 – 3.10). To help specification of conditional LGMs the following effects were hypothesized:

(a) antenatal CMD will have a significant direct effect on parameters of the LGM underlying the observed change in infant growth over the first 18 months of age (labelled A in figures 3.8-3.10),

(b) after adjusting for the direct effect of antenatal CMD on the parameters of LGMs, postnatal CMD will have a significant effect on the parameters of the underlying LGM of infant growth over the first 18 months of age (labelled B in figures 3.8-3.10),

(c) there will be significant continuity of CMD from antenatal to the two month postnatal period (labelled C in figures 3.8-3.10), and

(d) there will be a significant indirect effect of antenatal CMD on the parameters of the underlying LGM parameters through postnatal CMD (the product of B and C, i.e. C*B in figures 3.8-3.10).

After deciding on the best fitting unconditional LGM for each growth outcome only one of the hypothesized models fitting to the selected baseline model (i.e. one of figure 3.8-3.10) will be investigated.

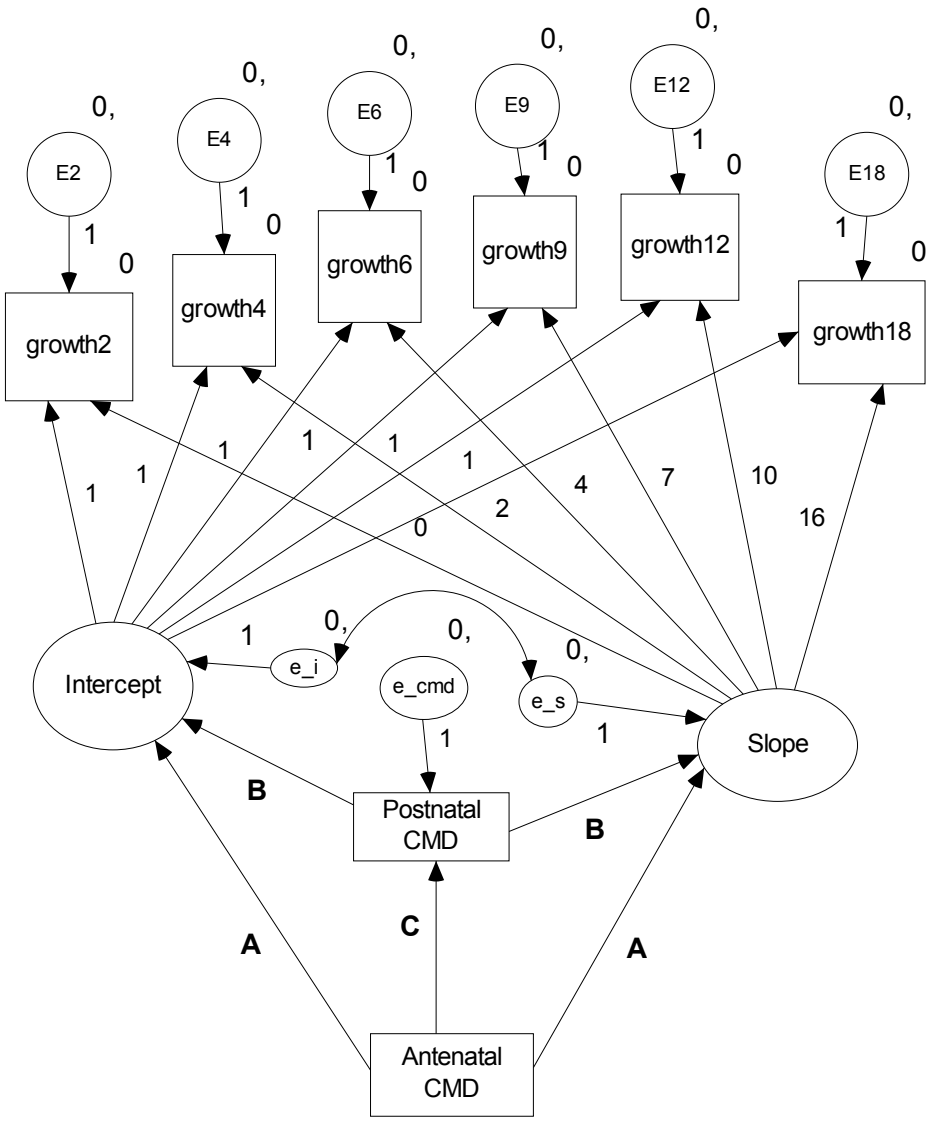

Figure 3.8: Hypothesized linear partially mediated LGM for infant growth in Butajira, Ethiopia

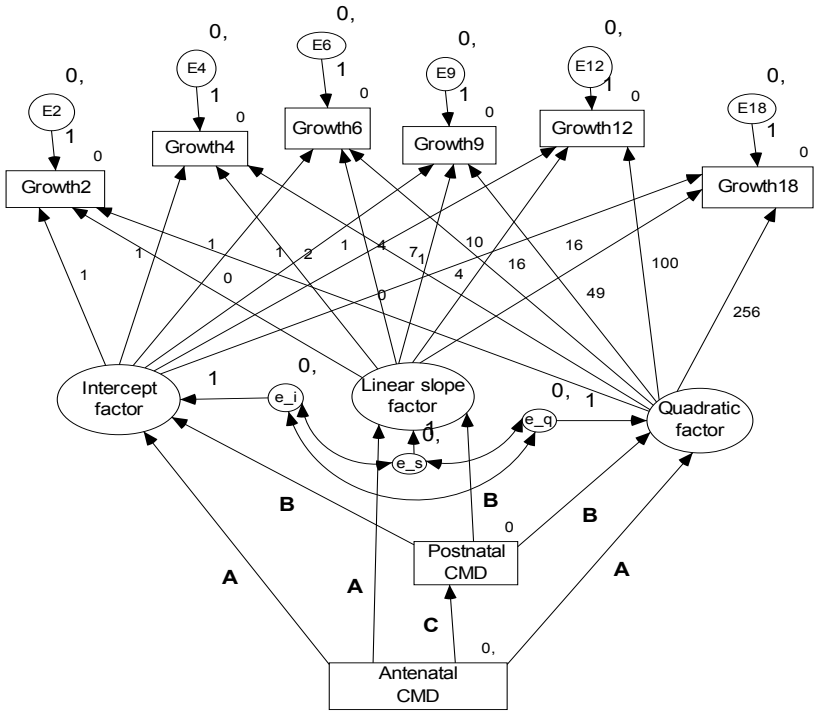

Figure 3.9: Hypothesized quadratic partially meditated LGM for infant growth in Butajira, Ethiopia

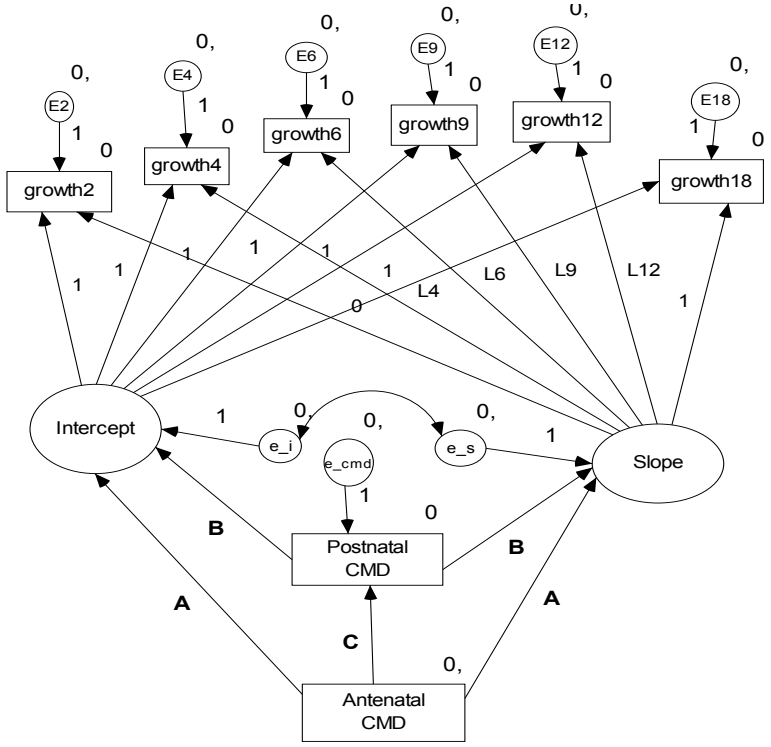

Figure 3.10: Hypothesized non-linear partially mediated LGM for infant growth in Butajira, Ethiopia

3.3.4.3.3 Mediating effects of birth weight and early infant feeding practices

In mediation LGMs antenatal and postnatal CMD are the main exposure variables. The interest is to evaluate the role of

(a) early infant feeding practices (i.e. withholding colostrum, delayed initiation of breast feeding, giving prelacteal feeding) and birth weight as mediators of the effect of antenatal CMD upon infant growth,

(b) the role of early infant illness (baby being severely ill and early infant diarrhoea) as mediators of the effect of persistent CMD on infant growth.

Figures 3.11-3.13 show the hypothesized relationship of CMD, mediating variables and growth outcomes. Only one of the three models is investigated depending on the best fitting level l LGM to be selected.

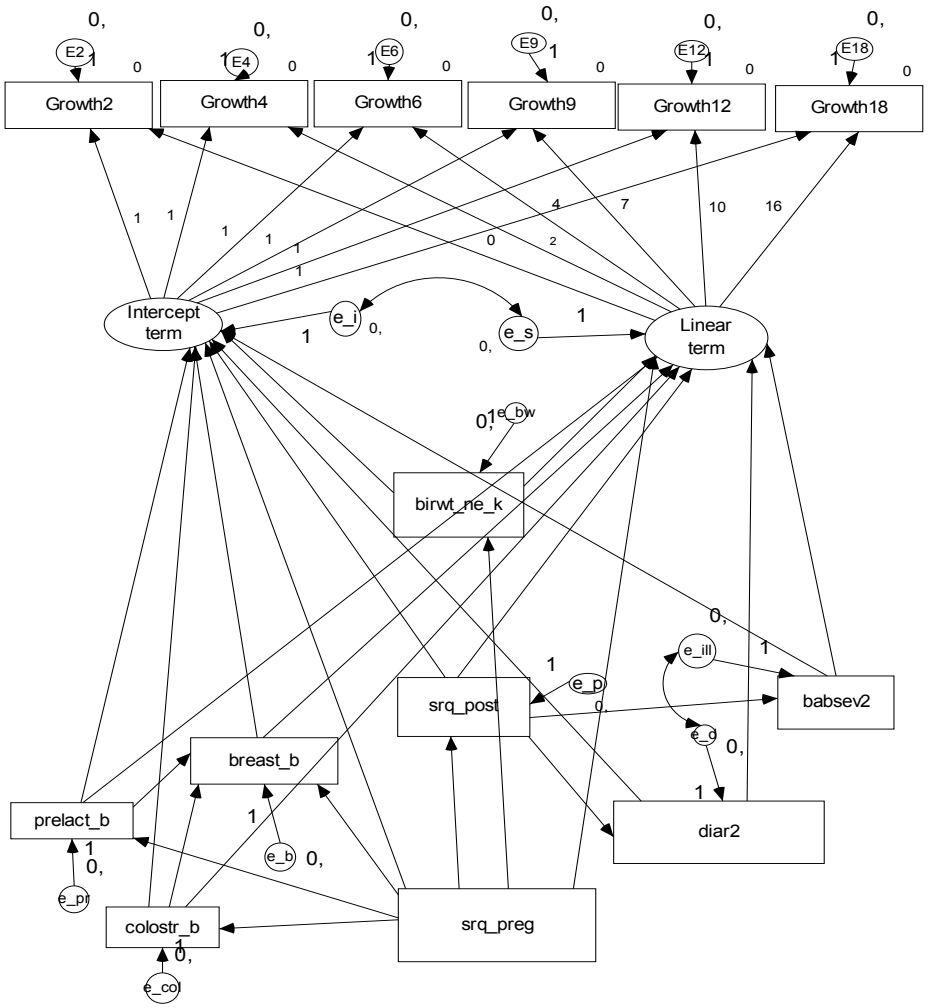

Figure 3.11: Hypothesized linear mediation LGM for infant growth in Butajira, Ethiopia

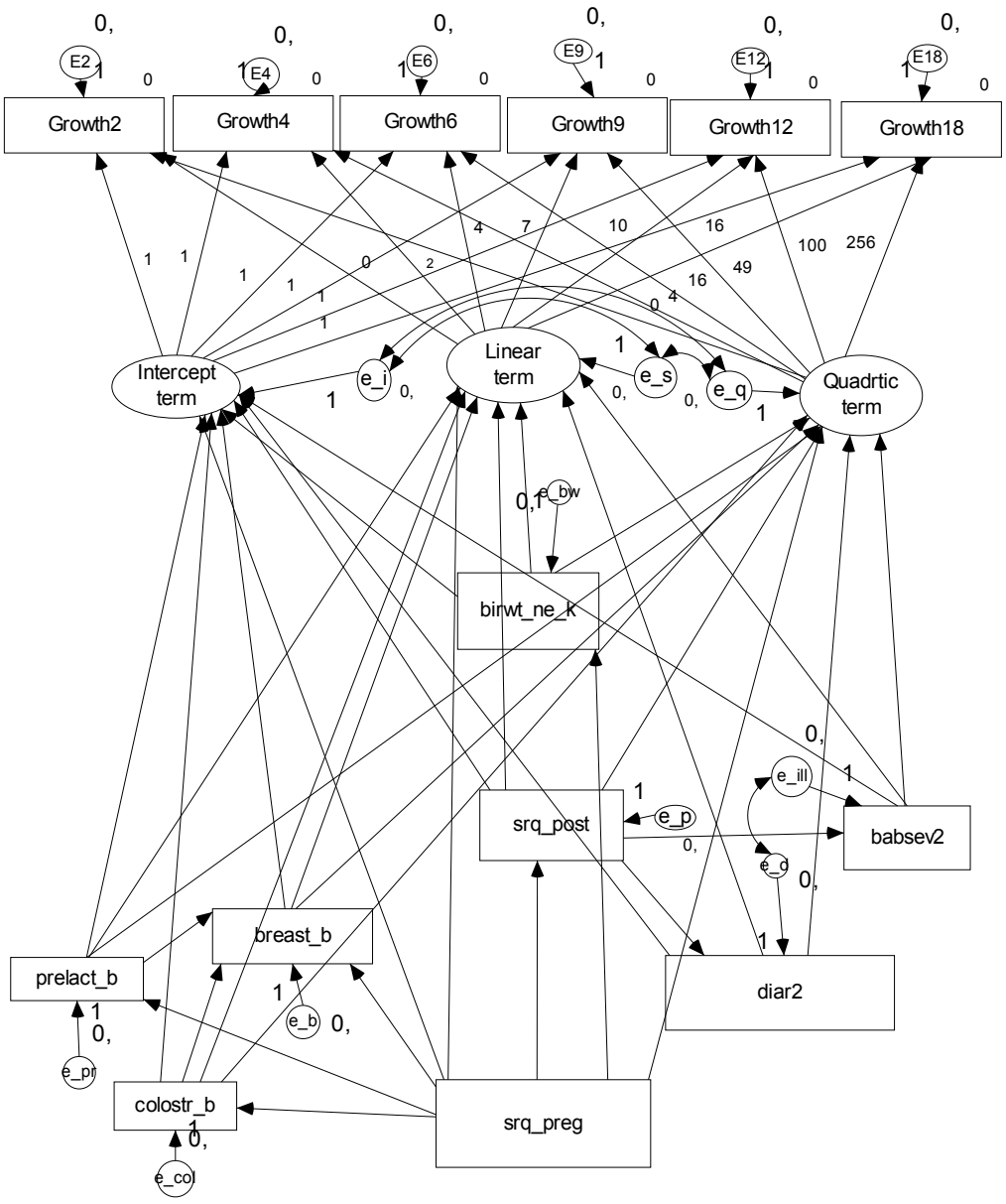

Figure 3.12: Hypothesized quadratic mediation LGM for infant growth in Butajira, Ethiopia

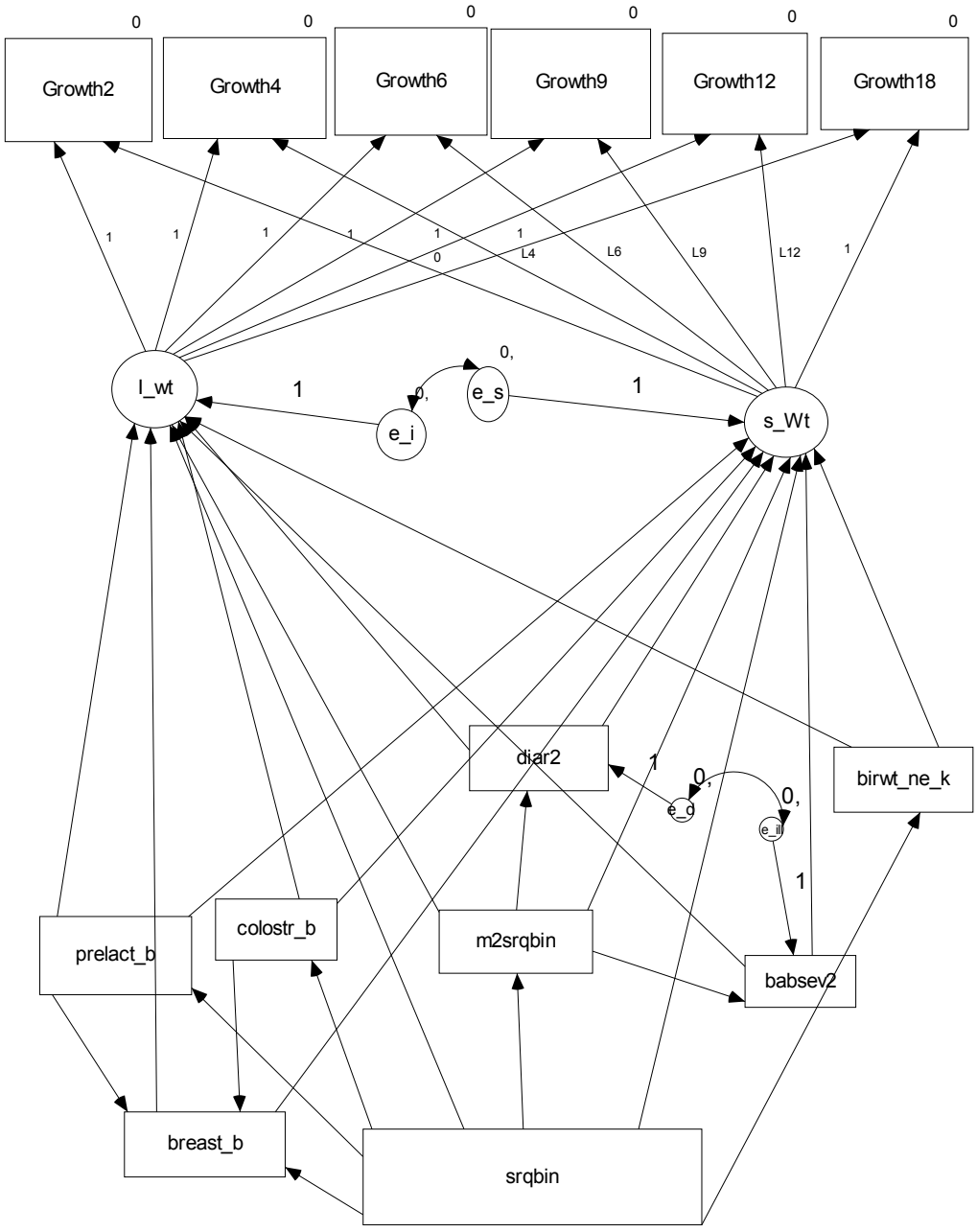

Figure 3.13: Hypothesized non-linear mediation LGM for infant growth in Butajira, Ethiopia

3.3.4.3.3 Fully adjusted effect of CMD on parameters of LGMs

Perinatal CMD is still the focus in the fully adjusted models. Conditional LGMs that simultaneously take account of antenatal CMD, postnatal CMD, mediating variables and pre-specified maternal, household and infant characteristics were specified as presented in figures 3.14 – 3.16. The only difference in the three hypothesized models is the way the level one growth (i.e. unconditional LGM for each growth outcome) was specified. In figure 3.14 the level one growth model is linear, in figure 3.15 the level one growth model is quadratic and in figure 3.16 the level one growth model is non-linear which is not quadratic. Based on the selection to be made after fitting the three candidate level one growth models only one of the hypothesized conditional LGM was investigated

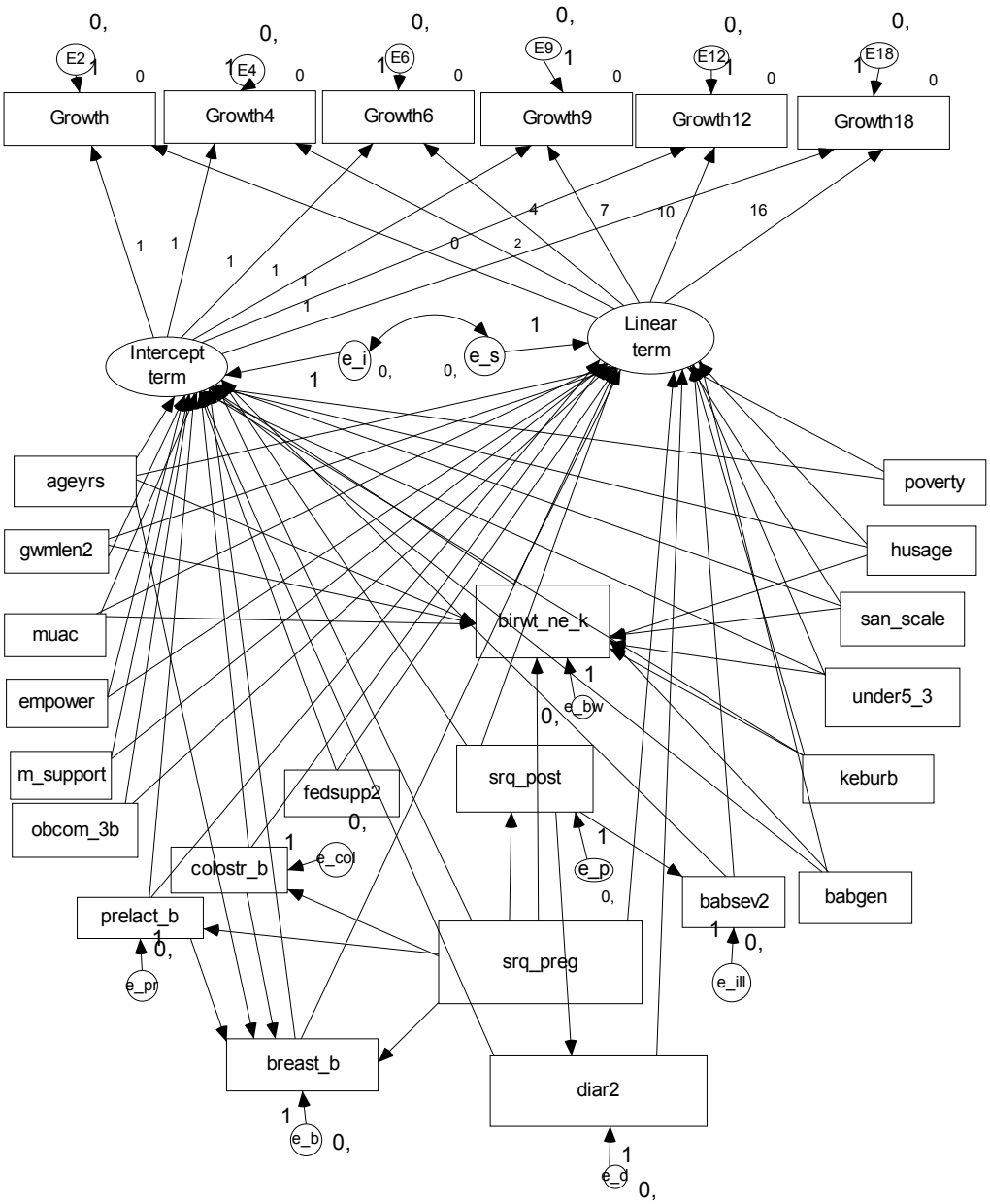

Figure 3.14: Hypothesized linear conditional LGM for infant growth in Butajira, Ethiopia

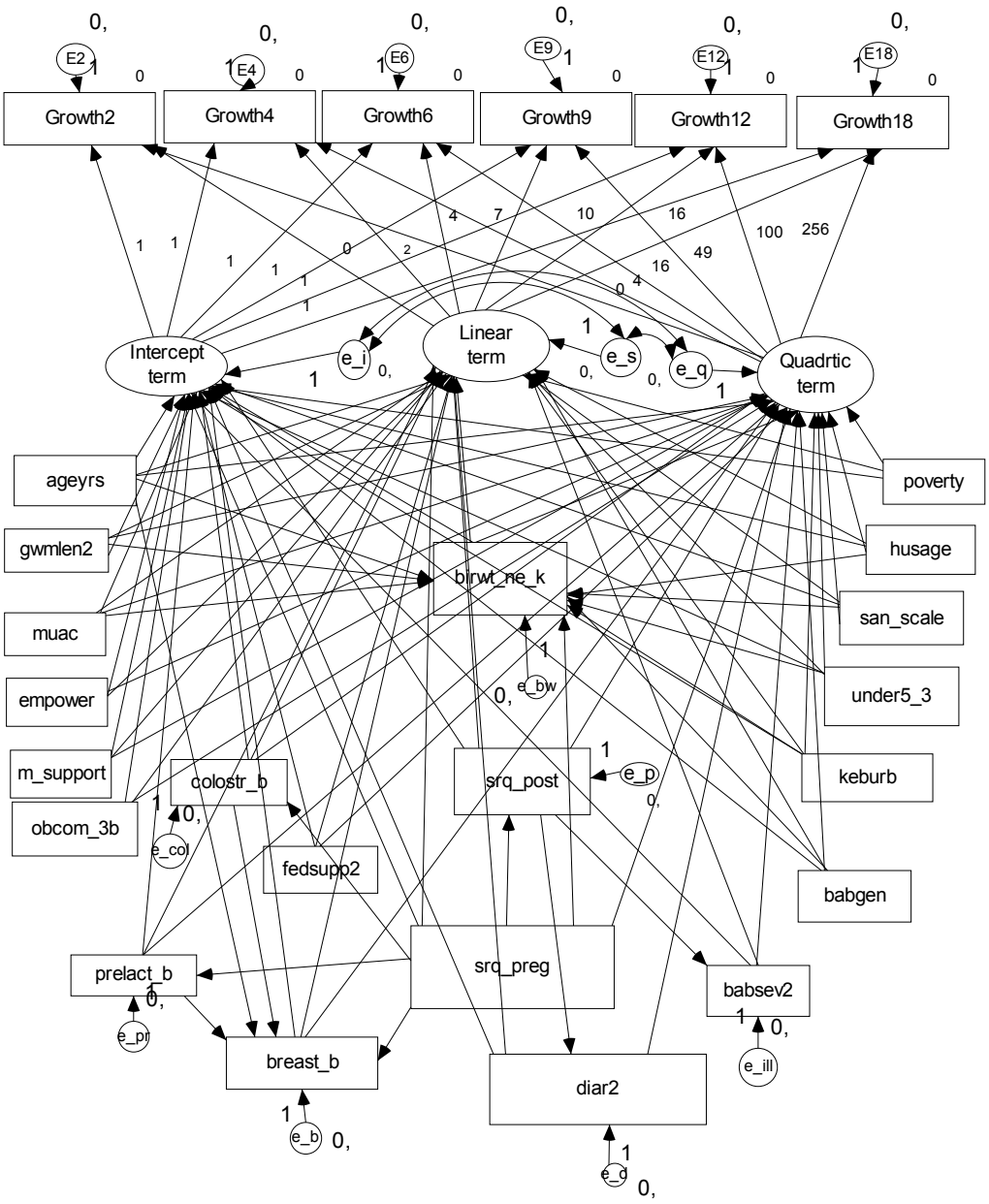

Figure 3.15: Hypothesized quadratic conditional LGM for infant growth in Butajira, Ethiopia

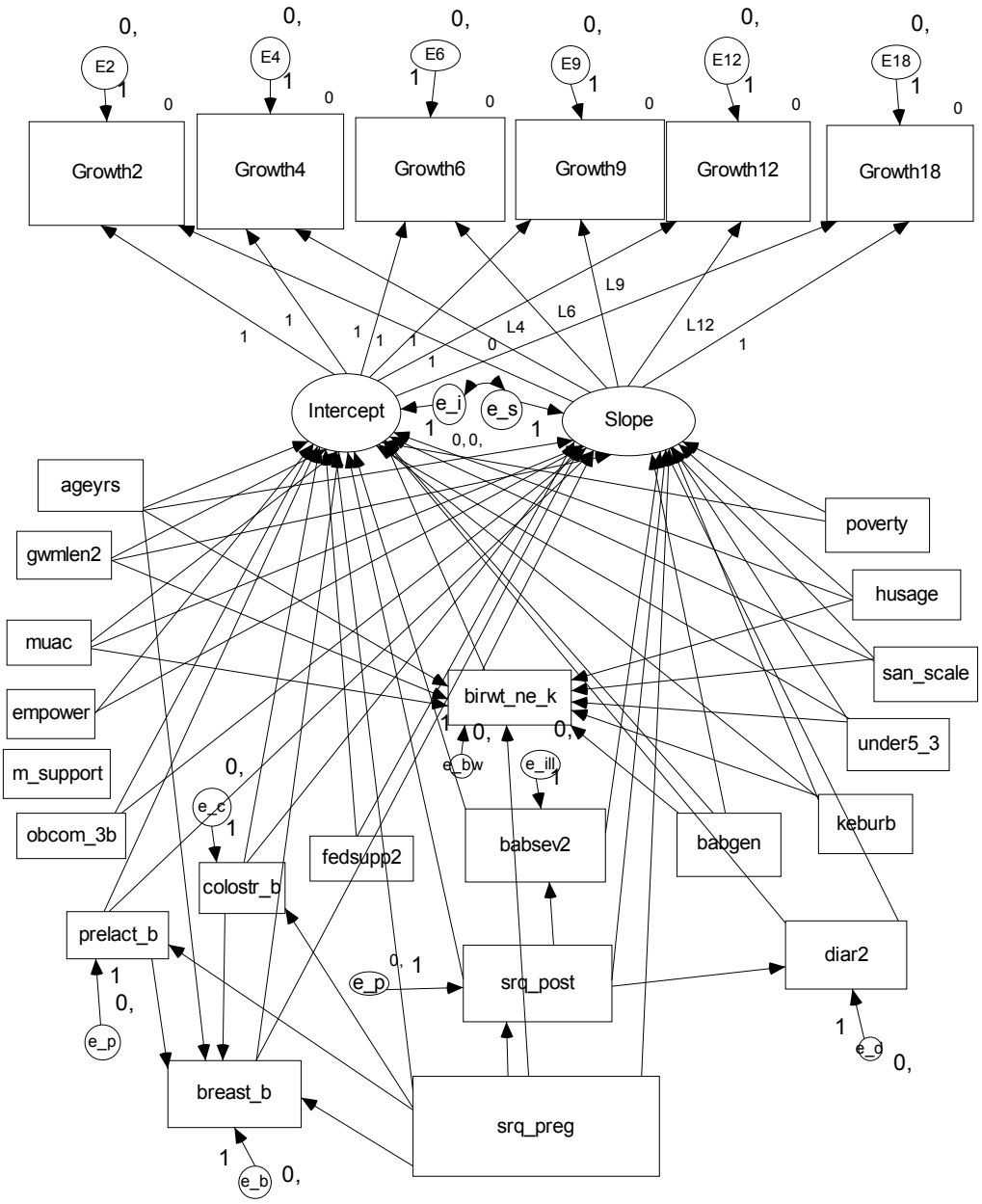

Figure 3.16: Hypothesized non-linear conditional LGM for infant growth in Butajira, Ethiopia

3.3.4.4 Assessing overall fit and individual parameters under different modelling strategies

In logistic regression, linear regression and multilevel growth modelling assessment of overall model goodness-of-fit is not the first mandatory step and the same tradition was followed in the presented data analysis. The Wald statistic was used to evaluate the statistical significance of individual regression coefficients, and residual plots to evaluate the degree of violations of regression assumptions as well as for identification of influential observations. In multilevel growth modelling theWald statistic was used to assess the statistical significance of individual parameters, residual plots to assess distributional assumptions of the random components, and deviance and information criteria to evaluate improvement of fit and decide on model improvement in case of nested models.

Following the same trend with the other modelling techniques Wald statistic was used in LGM to assess statistical significance of individual parameters. However, in order to come to the best fitting LGM, the first criterion was to evaluate the overall goodness-of-fit. Hence, **g**oodness-of-fit of different LGMs was assessed with the maximum-likelihood chi-square statistic, the comparative fit index (CFI)(Bentler 1990), Tuker-Lewis index (TLI) (Tucker and Lewis 1973), and the root mean squared error of approximation (RMSEA) (Steiger 1990; Hu and Bentler 1999). The values of CFI range from 0 to 1 and reflect the improvement in fit of a hypothesized model over a model of complete independence among the measured variables. Values approaching 0.95 or greater are desirable for the CFI. The RMSEA is a measure of fit per number of degrees of freedom, controlling for sample size; values less than .06 indicate a relatively good fit (Hu and Bentler 1999).

3.3.4.5 Handling of missing data under different modelling strategies

In logistic and linear regression the data were analyzed using all variables with complete records, and using list-wise deletion of cases with missing data. This is the default for all model fitting algorithms implemented in standard software that fits regression models for outcome variables from the family of generalized linear models. For multilevel modelling missing data were handled with the full information maximum likelihood

(FIML) missing data procedures available in Stata version 10 (StataCorp 2007). Similarly, for latent growth modelling missing data were handled with the full information maximum likelihood (FIML) missing data procedures available in AMOS version 7 (Arbuckle 1995-2006) and MPlus version 5 (Muthen and Muthen 1998-2009) for continuous and binary growth outcome variables, respectively. Using FIML rather than list-wise deletion allows for maximal use of available data, even from participants for whom information at certain time points is missing while attaining attractive properties of ML even with slight violation of assumptions (Schafer and Graham 2002).

CHAPTER 4: RESULTS - TRADITIONAL ANALYSIS

4.1 Introduction

In this chapter we will summarize results obtained from traditional data analysis methods which will be then compared with the findings from multilevel growth modelling and latent growth modelling to be presented in later chapters. To be specific this chapter will focus on results obtained from:

- descriptive summary of the cohort data,
- bivariate associations of pre-specified risk factors and to the six infant growth outcome variables (i.e. length, weight, length-for-age z, weight-for-age z, stunting and underweight) measured at 2, 6 and 12 months of infant age and
- results obtained from bivariate and multivariable linear regression and logistic regression taking each of the six infant growth outcomes at 2, 6 and 12 months of follow-up at a time.

We will have four sections in this chapter so that the content can be presented logically and easy to follow:

- Part I: In this section we will present a description of the cohort and bivariate associations of six infant growth measures with pre-specified risk factors of infant growth including maternal CMD.
- Part II: In this section of the chapter we will summarize unadjusted and adjusted results of logistic regression and linear regression where the outcome variables are the six infant growth measures at 2, 6 and 12 months and the independent variables are pre-specified household characteristics, maternal characteristic but not CMD, infant characteristics and early infant feeding practices of the mother.
- Part III: In this section we will summarize unadjusted and adjusted effects of maternal CMD on infant growth obtained from linear regression and logistic regression. Maternal CMD will be treated as three independent variables, namely, prevalence antenatal, prevalence postnatal and course of CMD from antenatal to postnatal (a categorical variable with four levels: negative at both time points, remitted antenatal, incident postnatal and chronic)
- Part IV: In this section the main findings will be summarized

4.2 Description of the cohort and bivariate associations of infant growth with pre-specified risk factors

4.2.1 Description of the cohort

Cohort characteristics

Initial recruitment and detailed description of follow-up is presented in figure 4.1. From the cohort of 1065 women 1006 give birth to singleton live babies and 966 of them were eligible for anthropometric growth follow-up starting at 2 months of age. We have now analysed anthropometric growth data from 912 singleton infants at two months of age, 888 at 6 months of age, 880 at 9 month of age, 926 at 12 months of age and 912 infants at 18 months of age. The main causes of attrition were stillbirth (n=40), neonatal mortality (n=35), and multiple births (15 twins and 1 triple birth). Temporary (n=95) and permanent (n=10) out-migration also contributed to loss to follow-up. Nineteen cases at six months, 14 cases at twelve months and 6 cases at 18 months were excluded due to measurement error on the outcome and flagged by the Anthro software (Monika, Elaine et al. 2006) as being outliers during standardisation of growth measurements.

Selected background characteristics of study participants at baseline, two, six and twelve months follow-up are summarized in table 4.1. At baseline 99.0% of the women were married and on average younger (mean age 27.9 years, sd =6.4 years) than their husbands (mean age 36.2 years, sd = 9.2 years) (Table 4.1). The majority of women were non-literate (85%), reported their occupation as housewives or farming (88%), and belonged to the Meskan ethnic group (46%). Comparatively, the non-literacy rate was lower among the husbands (31%). Most of the women reported having protected water (71%) and a toilet facility (63%) but few of them disposed of rubbish in a sanitary way (22%). During the study period there was no significant change in the composition of the study participants in terms of major maternal socio-demographic and household characteristics.

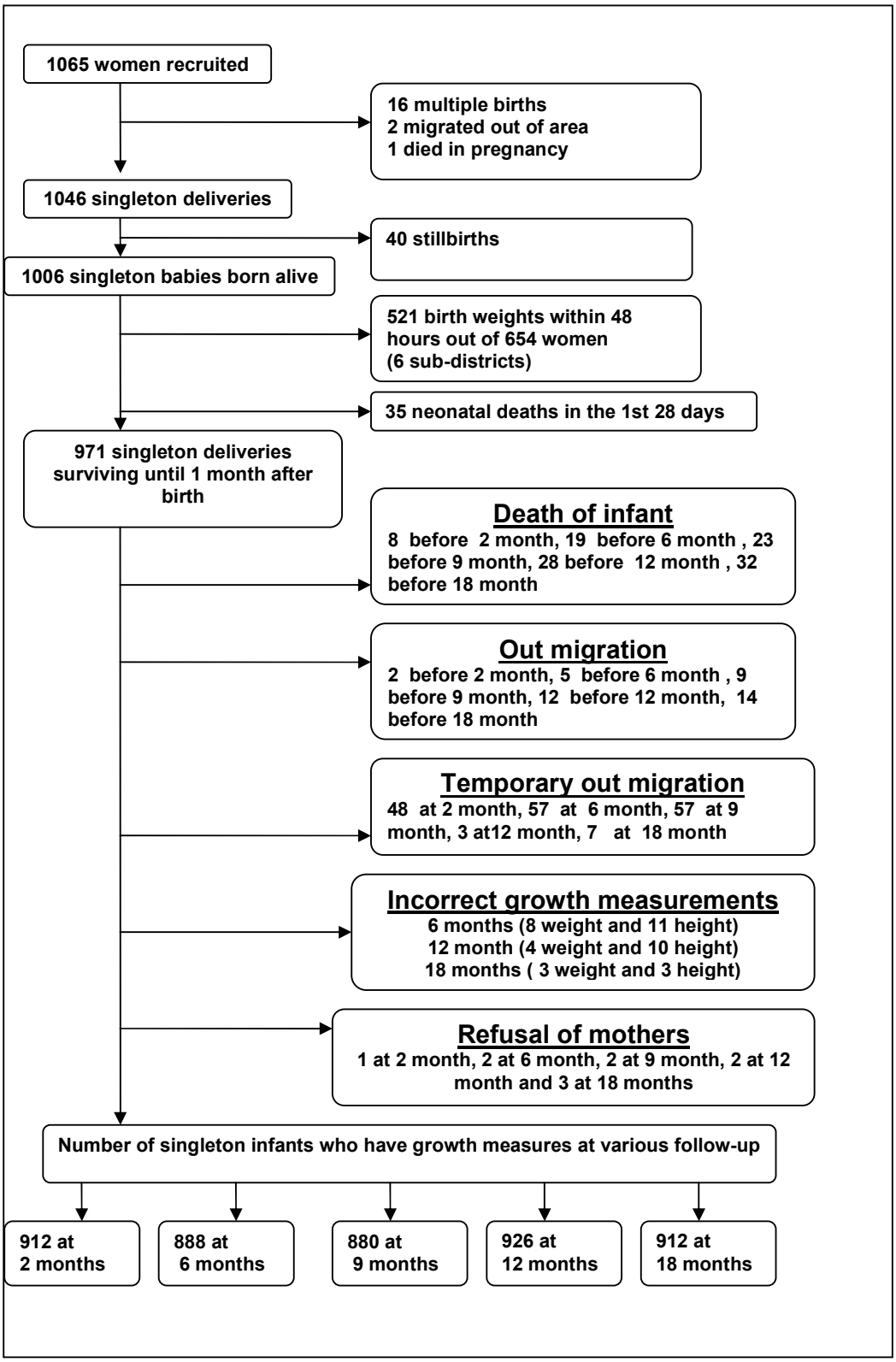

Figure 4.1: Follow-up of the P-MaMiE cohort from screening up to one year postnatal

Table 4.1: Selected background characteristics of P-MaMiE study participants [Number(%) or Mean(SD)]

Characteristics		follow-up time		
	baseline	2 month	6 month	12 month
Maternal Characteristics				
Religion: Muslim	826(77.6)	694(77.0)	674 (77.2)	722 (78.0)
Orthodox Christian	161(15.1)	140(15.5)	133 (15.3)	139 (15.0)
Protestant	66(6.2)	57(6.3)	56 (6.4)	55 (5.9)
Catholic	12(1.1)	10(1.1)	10 (1.2)	10 (1.1)
Ethnicity: Meskan	485(45.5)	413(45.8)	404 (46.3)	436 (47.1)
Mareko	147(13.8)	119(13.2)	119 (13.7)	123 (13.3)
Silti	257(24.1)	212(23.5)	200 (22.8)	218 (23.5)
Sodo	85(8.0)	74(8.2)	69 (7.9)	68 (7.3)
Others	91(8.5)	83(9.2)	81 (9.3)	81 (8.8)
Currently married	1055(99.1)	892(99.0)	865 (99.1)	917 (99.0)
Occupation				
Housewife or farming	933(87.9)	785(87.3)	759 (87.1)	805 (87.1)
Self or paid employee	129(12.2)	114(12.7)	112 (12.9)	119 (12.9)
Age (years)	26.9(6.4)	26.9(6.2)	26.8(6.2)	26.9(6.2)
Educational status				
Literate	165(15.5)	139(15.4)	134 (15.4)	141 (15.2)
Non-literate	900(84.5)	762(84.6)	739 (84.7)	785 (84.8)
Household characteristics				
Age of husband (years)	36.2(9.2)	36.0(8.7)	36.0(8.8)	36.1(8.9)
Educational status of husband				
Literate	726(68.6)	620(69.3)	594 (68.4)	632 (68.7)
Non-literate	333(31.4)	275(30.7)	274(31.6)	288 (31.3)
Main source of water				
Protected supply	752(70.8)	627(69.7)	601 (69.0)	644 (69.7)
Unprotected supply	310(29.2)	272(30.3)	270 (31.0)	280 (30.3)
Sanitary condition				
Have toilet facilities	674(63.3)	572(63.5)	553 (63.3)	582 (62.9)
No proper toilet facilities	391(36.7)	329(36.5)	320 (36.7)	344 (37.2)
Rubbish disposal				
Buries, burns or others	238(22.4)	202(22.4)	192 (22.0)	200 (21.6)
Disposes on field	826(77.6)	698(77.6)	680 (78.0)	725 (78.4)
Traditional surgical practices before two months of age				
Uvulectomy: Performed		16(1.8)	16 (1.8)	16 (1.7)
Not Performed		880(98.2)	853 (98.2)	905 (98.3)
Circumcision of girls:				
Performed		7(1.6)	6 (1.4)	7 (1.5)
Not Performed		437(98.4)	415 (98.6)	447 (98.5)
Circumcision of boys:				
Performed		192(42.4)	187 (41.6)	196 (41.9)
Not Performed		261(57.6)	262 (58.4)	272 (58.1)

The prevalence of regular khat chewing during pregnancy (28%) was significantly higher than during the two month postnatal period (21%) (OR=1.84; 95% CI: 1.38, 2.47). At one year follow-up 51.5% of the study participants were current users of khat and 24.7% were chewing khat at least once per week. The prevalence of regular alcohol

consumption was generally low among the study participants and reduced significantly during the two month postnatal period compared to pregnancy (1.6% & 5.1%; $p < 0.001$). At one year follow-up 1.0% of the women were drinking alcohol one to two times a week and 2.5% were drinking alcohol one to three times per month. In this cohort, harmful traditional practices were relatively rare with a low prevalence of performing uvulectomy (1.8% before two months of age and 4.8% before one year of age), female circumcision (1.5%) before two months of age or at one year follow-up and milk teeth extraction (3.9%) in the first year of infancy.

Fifty four percent of women attended antenatal care, 90% delivered at home and only 24% of deliveries were attended by trained personnel at any level. Despite this, a high proportion of mothers reported giving colostrum to their newborn (82%), 32% initiated breastfeeding within the first hour and 16% reported giving pre-lacteal food to their new born. At the age of two months 59% of the infants had already received at least one type of vaccination: 53% Polio, 29% DPT, and 35% BCG.

At 12 and at 18 months of age 99.6% and 96.1% of the infants were breastfeeding, respectively. At the six month anthropometric assessment 33.4% of the mothers reported that their infants were currently ill; this prevalence was reduced to 30.6% during twelve month growth assessment. During the same follow-up 95.2% of the mothers reported at least one infant illness episode since birth, 60.0% of whom had thought that their baby was going to die because of the severity of the episode. Introduction of supplementary feeding occurred for 33.8% of infants by six months, 36.8% during 7-9 months and 28.3% after 9 months.

Prevalence of infant undernutrition

The prevalence of infant stunting increased throughout the first 18 months of follow-up, prevalence of underweight fell after nine months of age and prevalence of wasting was decreasing starting at two months of age (Figure 4.2).

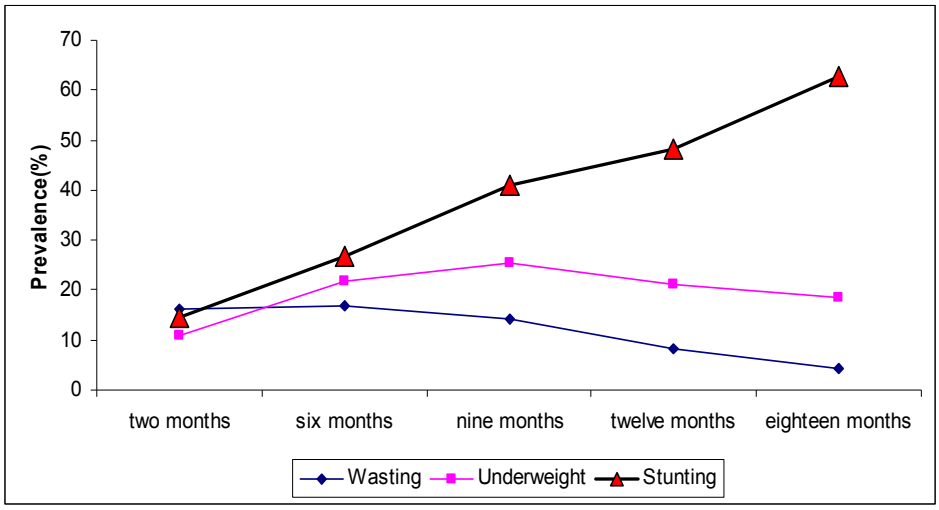

Figure 4.2: Prevalence of infant undernutrition from two to eighteen months of age

The overall prevalence of stunting and underweight were 14.6% and 10.8% at the age of two months, 26.7% and 21.7% at the age of six months and 48.1% and 21.2% at the age of twelve months, respectively.

Comparison of the cohort with 2006 WHO child growth standards

Mean length-for-age z score had negative trend between two and eighteen months, and the rate was relatively steeper before nine months. Mean weight-for-age z of infants started to improve after nine months of age (figure 4.3).

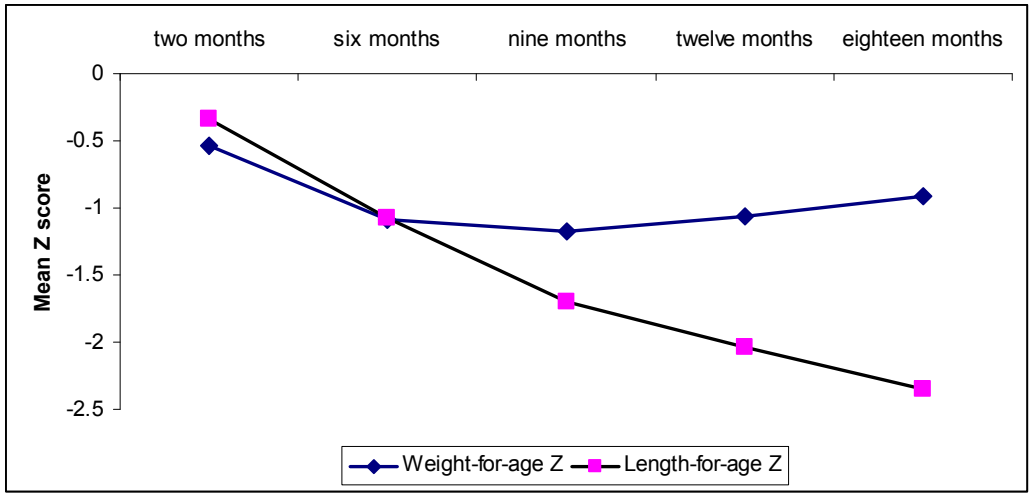

Figure 4.3: Weight and length of infants standardized using 2006 WHO reference population

Compared to the 2006 WHO child growth standards, which have a mean of zero and standard deviation of one, infants in this cohort in the first 18 months of life were significantly lighter as reflected in their mean (95% CI) weight-for-age z score at two months (-0.54; -0.62 to -0.47), six months (-1.09; -1.17 to -1.00), nine months(-1.17; -1.26, -1.08), twelve months (-1.06; -1.14 to -0.98), and eighteen months (-0.91; -0.99, -0.82). They were also shorter as reflected in their mean (95% CI) height-for-age z score at two months (-0.34; -0.45 to -0.23), six months (-1.08; -1.19 to -0.98), nine months (-1.70; -1.80, -1.60), twelve months of age (-2.04; -2.14 to -1.94), and eighteen months (-2.35; -2.43 to -2.26).

Weight and length of infants from two to eighteen months of age

At the age of two months infants attained a mean weight of 5.3kg with standard deviation of 0.9kg and a mean length of 57.7 cm with standard deviation of 4.3 cm. At the age of eighteen months their mean weight was 9.6 kg with standard deviation of 1.5kg (i.e. an increase of 4.3kg in 16 months) and their mean length was 74.8 kg with standard deviation of 3.8cm (i.e. an increase of 17.1cm in 16 months). As it would be expected the trend of change in mean weight and mean length were consistently increasing over the study period (Figures 4.4 and 4.5).

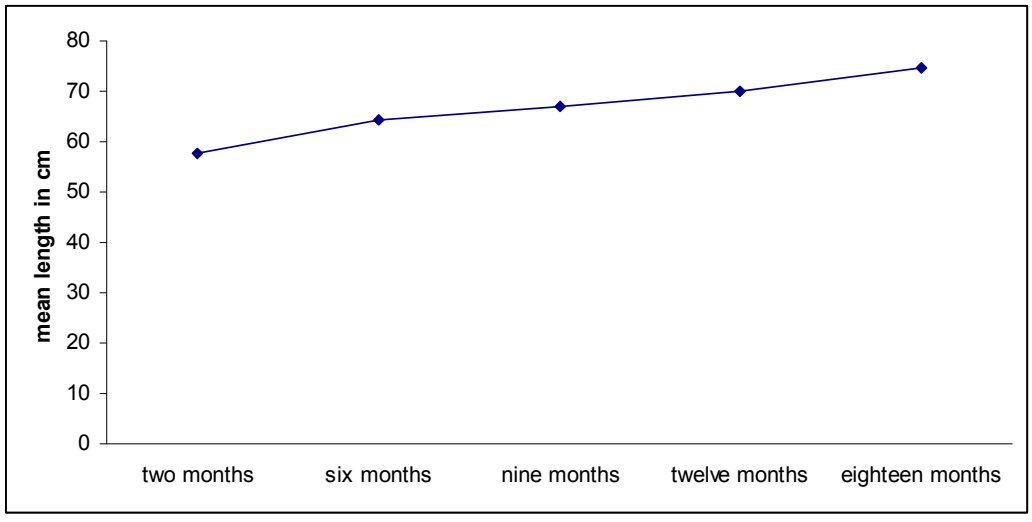

Figure 4.4: Mean length of infants at various follow-up time points

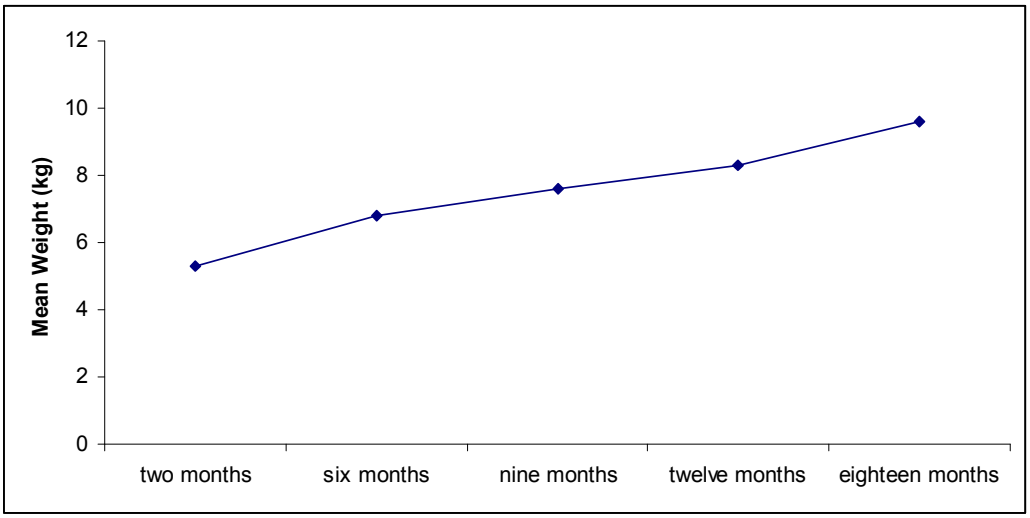

Figure 4.5: Mean weight of infants at various follow-up time points

Prevalence of maternal CMD

Of the total 1065 women recruited in the third trimester of pregnancy 128 (12.0%) had antenatal CMD. One thousand and forty-five of the mothers (98.1%) were re-interviewed at two months post partum and 56 (5.4%) had postnatal CMD including 26 (2.8%) incident cases. Among one thousand and thirty seven women who were interviewed at one year follow-up 6.8% had high level of CMD (i.e. 4.3% among those who had no high level of CMD during recruitment as well as at two months post partum follow-up, 13.0% among whose who had high level of CMD at pregnancy and remitted at two month post partum, 23.1% among incident cases during the two month post partum follow-up and 43.3% among women who had high level of CMD at recruitment and at two month post partum).

4.2.2 Bivariate associations of infant growth with pre-specified risk factors

Length

In this cohort prenatal maternal CMD was not significantly associated with mean infant length at the ages of 2, 6 and 12 months (table 4.2). However, among two month old infants there was a trend of having a shorter mean length if their mothers had high level of symptoms during a two month postnatal follow-up (i.e. prevalent or incident postnatal)

Table 4.2: Bivariate associations of perinatal maternal CMD and infant length in cm at 2, 6 and 12 months of age

Timing of CMD measure and its categories	*2 months Mean(SE) length*	*6 months Mean(SE) length*	*12 months Mean(SE) length*
Pregnancy			
Low SRQ score	57.64(0.16)	64.26(0.13)	69.98(0.13)
High SRQ score	58.25(0.37)	64.45(0.30)	70.00(0.42)
P-value*	0.17	0.61	0.97
Two month postnatal			
Low SRQ score	57.75(0.15)	64.25(0.12)	69.97(0.13)
High SRQ score	56.71(0.60)	64.81(0.47)	70.10(0.57)
p-value*	0.13	0.32	0.83
Pregnancy or postnatal			
Low SRQ at all time point	57.66(0.16)	64.24(0.13)	70.0(0.13)
High SRQ score at both time points	56.87(0.73)	64.95(0.65)	70.5(0.93)
High SRQ score at Postnatal only	56.50(1.03)	64.58(0.67)	69.6(0.54)
High SRQ score at Pregnancy only	58.65(0.42)	64.29(0.34)	69.9(0.46)
p-value*	0.10	0.77	0.87

SE: stanadard error

Bivariate associations of infant length and selected background characteristics are summarized in table 4.3. Female gender and low birth weight were consistently associated with reduced infant length across the three time points and urban residence was significantly associated with an increase in length at two and twelve months but not at six months of age. Having a toilet facility was significantly associated with an increase in length at 6 and 12 months. Compared to their counterparts infants whose fathers were literate, families burn rubbish and use water from protected source were significantly longer in their mean length at one year of age. At the age of two months infants had significantly shorted mean length if their mothers were non-literate or if their families were users of water from protected source.

Table 4.3: Mean(SD) of length of infants at two, six and twelve months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		12 month follow-up	
	Mean(SD)	p-value	Mean(SD)	p-value	Mean(SD)	p-value
Sex						
Male	58.16(4.62)	0.002	64.64(3.55)	0.002	70.36(3.81)	0.002
Female	57.25(4.00)		63.91(3.34)		69.59(3.69)	
Birth weight						
Normal	64.63(3.35)	<0.001	64.63(3.35)	0.001	70.30(3.70)	0.001
Low	63.45(3.30)		63.45(3.30)		68.60(4.20)	
Missing	63.77(3.62)		63.77(3.62)		69.50(3.80)	
Residence						
Rural	57.54(4.48)	0.003	64.21(3.41)	0.142	69.77(3.64)	0.000
Urban	58.74(3.27)		64.71(3.79)		71.36(4.32)	
Under five children						
None	57.82(3.70)	0.360	64.21(3.48)	0.951	70.40(3.90)	0.201
Only one	57.85(4.17)		64.30(3.55)		69.90(3.80)	
Two or more	57.39(5.01)		64.30(3.30)		69.80(3.50)	
Sanitation facility						
Toilet available	57.60(4.56)	0.301	64.63(3.40)	0.000	70.26(3.90)	0.004
Open field	57.91(3.95)		63.70(3.50)		69.52(3.51)	
Rubbish disposal						
Burn	58.23(3.82)	0.052	64.53(3.56)	0.264	70.61(3.90)	0.008
Open field	57.56(4.48)		64.21(3.44)		69.81(3.72)	
Water source						
Protected	57.40(4.16)	0.001	64.22(3.58)	0.423	70.29(3.76)	0.000
Unprotected	58.42(4.70)		64.42(3.20)		69.26(3.72)	
Initiation of breast feeding						
<= 1 hour	58.09(3.78)	0.098	64.47(3.42)	0.242	69.86(3.89)	0.495
>1 hour	57.57(4.52)		64.18(3.49)		70.04(3.75)	
Colostrums						
Given	57.71(4.23)	0.616	64.22(3.47)	0.292	69.93(3.79)	0.414
Denied	57.90(4.59)		64.54(3.46)		70.20(3.66)	
Pre-lacteal feeding						
Given	57.57(3.69)	0.803	64.21(3.09)	0.899	69.69(3.26)	0.603
Not given	57.75(4.33)		64.28(3.49)		70.00(3.77)	
Breastfeeding at two months						
Exclusive	57.74(4.36)	0.520	64.23(3.40)	0.428	69.91(3.70)	0.252
Non-exclusive	57.49(4.32)		64.48(3.71)		70.30(4.07)	
Maternal education						
Literate	58.51(3.80)	0.005	64.24(4.28)	0.838	70.46(4.43)	0.058
Non-literate	57.51(4.45)		64.30(3.23)		69.87(3.58)	
Father's education						
Literate	57.79(4.41)	0.404	64.34(3.46)	0.447	70.19(3.78)	0.021
Non-literate	57.53(4.21)		64.15(3.49)		69.57(3.74)	

SD: Standard deviation

None of the correlation coefficients of infant length at various follow-up time points and selected household characteristics summarized in table 4.4 is very large although maternal MUAC was positively correlated with infant length at six and twelve months and maternal age, poverty index and poor sanitary condition scale were inversely correlated with infant length at twelve months of age.

Table 4.4: Correlation of selected background parental characteristics and infant length at two, six and 12 months of age

	Length					
Background characteristics	**Two months**		**Six months**		**Twelve months**	
	Correlation	**p-value**	**Correlation**	**p-value**	**Correlation**	**p-value**
Maternal age	0.019	0.565	0.024	0.477	-0.067	0.042
Age of father	0.002	0.951	0.016	0.644	-0.039	0.236
Maternal MUAC	0.042	0.204	0.113	0.000	0.071	0.031
Poverty index	-0.058	0.084	0.011	0.750	-0.120	0.000
Sanitary index	0.053	0.108	-0.066	0.050	-0.153	0.000
Maternal support	0.016	0.638	-0.014	0.673	-0.053	0.111
Maternal autonomy	-0.027	0.415	-0.034	0.306	0.040	0.224

Weight: Bivariate associations of prenatal maternal CMD and infant weight at two, six and 12 months of age are summarized in table 4.5. Contrary to our expectation prevalent postnatal CMD was significantly associated with increased mean weight at two and six months of age which was no longer significant at twelve months of age. Although statistically non-significant infants whose mothers had chronic CMD from pregnancy to two months postnatal had a trend in their mean weight of being heavier at two, six and 12 months of age compared to infants whose mothers had no CMD in either of the two time points or had it only at one of the two time points.

Table 4.5: Bivariate associations of prenatal maternal CMD and infant weight in kg at 2, 6 and 12 months of age

Timing of CMD measure and its categories	*2 months Mean(SE)*	*6 months Mean(SE)*	*12 months Mean(SE)*
Pregnancy			
Low SRQ score	5.25(0.03)	6.78(0.04)	8.35(0.04)
High SRQ score	5.37(0.09)	6.75(0.12)	8.30(0.13)
P-value*	0.16	0.80	0.72
Two month postnatal			
Low SRQ score	5.25(0.03)	6.76(0.04)	8.33(0.04)
High SRQ score	5.51(0.14)	7.14(0.15)	8.55(0.14)
p-value*	0.06	0.02	0.26
Pregnancy or postnatal			
Low SRQ at all time point	5.24(0.03)	6.77(0.04)	8.35(0.04)
High SRQ score at both time points	5.55(0.21)	7.17(0.23)	8.87(0.20)
High SRQ score at Postnatal only	5.46(0.20)	7.10(0.15)	8.12(0.16)
High SRQ score at Pregnancy only	5.32(0.10)	6.62(0.13)	8.14(0.16)
p-value*	0.23	0.09	0.07

SE: Standard error

Table 4.6: Mean(SD) of weight of infants at two, six and twelve months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		*12 month follow-up*	
	Mean(SD)	p-value	Mean(SD)	p-value	Mean(SD)	p-value
Sex						
Male	5.42(0.78)	0.000	6.93(1.10)	0.000	8.50(1.27)	0.000
Female	5.10(0.82)		6.60(0.91)		8.18(1.21)	
Birth weight						
Normal	6.79(1.03)	0.000	6.79(1.03)	0.000	8.40(1.26)	0.001
Low	6.05(0.90)		6.05(0.90)		7.64(1.25)	
Missing	6.85(1.00)		6.85(1.00)		8.32(1.21)	
Residence						
Rural	5.26(0.86)	0.738	6.69(0.99)	0.000	8.29(1.26)	0.001
Urban	5.28(0.90)		7.30(1.07)		8.69(1.13)	
Under five children						
None	5.20(0.89)	0.483	6.78(1.09)	0.942	8.49(1.11)	0.123
Only one	5.29(0.86)		6.78(1.04)		8.34(1.30)	
Two or more	5.26(0.86)		6.75(0.96)		8.24(1.25)	
Sanitation facility						
Toilet available	5.28(0.88)	0.323	6.91(1.00)	0.000	8.42(1.28)	0.016
Open field	5.22(0.85)		6.54(1.03)		8.21(1.20)	
Rubbish disposal						
Burn	5.34(0.88)	0.127	7.09(0.96)	0.000	8.47(1.14)	0.113
Open field	5.24(0.86)		6.68(1.03)		8.31(1.28)	
Water source						
Protected	5.29(0.88)	0.196	6.89(1.02)	0.000	8.49(1.19)	0.000
unprotected	5.21(0.83)		6.52(1.00)		8.00(1.33)	
Initiation of breast feeding						
<= 1 hour	5.37(0.89)	0.014	6.94(1.10)	0.001	8.46(1.35)	0.057
>1 hour	5.21(0.85)		6.69(0.99)		8.29(1.20)	
Colostrums						
Given	5.29(0.89)	0.071	6.80(1.04)	0.111	8.38(1.26)	0.049
Denied	5.15(0.75)		6.65(0.99)		8.17(1.19)	
Pre-lacteal feeding						
Given	5.23(0.70)	0.835	6.68(1.09)	0.552	8.20(1.21)	0.448
Not given	5.26(0.87)		6.78(1.03)		8.35(1.25)	
Breastfeeding at two months						
Exclusive	5.26(0.85)	0.821	6.75(1.00)	0.203	8.32(1.28)	0.306
Non-exclusive	5.25(0.95)		6.87(1.13)		8.44(1.12)	
Maternal education						
Literate	5.41(0.97)	0.011	7.00(1.18)	0.001	8.53(1.23)	0.019
Non-literate	5.23(0.84)		6.72(0.98)		8.29(1.25)	
Father's education						
Literate	5.28(0.85)	0.425	6.85(1.01)	0.002	8.42(1.26)	0.009
Non-literate	5.23(0.90)		6.61(1.04)		8.18(1.23)	

SD: Standard deviation

Bivariate associations summarized in table 4.6 show that infants who are male, born with normal birth weight, initiated breast feeding not later than one hour after birth, and had a literate mother performed significantly better in their weight at all time points although the statistical significance of the effect of delayed initiation of breast feeding at one year follow-up was marginal. Similarly, infants who had a literate father and whose parents lived in the urban area, had a toilet facility, and used water from a protected source

performed significantly better in their weight at six and twelve months of age. At the age of six months having a family which burns its rubbish rather than discarding it on an open field was associated with increased mean weight. Maternal MUAC was significantly and positively correlated with infant weight at all time points but parental age was significantly and inversely associated with infant weight at six months of age (i.e. age of the father) and at twelve months of age (i.e. age of the mother) (Table 4.7). Higher score on poverty index and poor sanitary condition scale were both significantly and inversely correlated with infant weight at the age of six and twelve months. However, none of these correlation coefficients were large showing low percentage of variability in the outcome explained by each of the background characteristics

Table 4.7: Correlation of selected background parental characteristics and infant weight at selected follow-up tome points

	Weight					
Background	**Two months**		**Six months**		**Twelve months**	
characteristics	**Correlation**	**p-value**	**Correlation**	**p-value**	**Correlation**	**p-value**
Maternal age	0.021	0.518	-0.044	0.195	-0.149	0.000
Age of father	-0.003	0.918	-0.071	0.037	-0.117	0.000
Maternal MUAC	0.111	0.001	0.169	0.000	0.120	0.000
Poverty index	-0.042	0.208	-0.191	0.000	-0.130	0.000
Sanitary index	-0.065	0.050	-0.244	0.000	-0.155	0.000
Maternal support	-0.020	0.548	-0.025	0.470	-0.036	0.281
Maternal autonomy	-0.025	0.445	0.034	0.311	0.014	0.665

Length-for-age

There was a statistically non-significant overall trend that showed high level of CMD at base line and at two month postnatal period being associated with smaller mean length-for-age z (table 4.8). When the course of CMD between the two time points were considered relatively larger negative effect was due to incident postnatal followed by chronic CMD. At the age of two months mean length-for-age z was significantly associated with prevalent postnatal CMD and course of CMD (i.e. a four category CMD variable).

Maternal MUAC was positively and significantly associated with length-for-age z at 6 and 12 months of age. Higher score on poverty index or poor sanitary condition scale were significantly and inversely associated with length-for-age z (t**able 4.9**)

Table 4.8: Bivariate associations of maternal CMD and infant length-for-age z at two, six and 12 months of age

Timing of CMD measure and its categories	*2 months Mean(SE)*	*6 months Mean(SE)*	*12 months Mean(SE)*
Pregnancy			
Low SRQ score	-0.36(0.06)	-1.07(0.06)	-2.03(0.05)
High SRQ score	-0.19(0.15)	-1.17(0.14)	-2.08(0.17)
P-value*	0.33	0.55	0.74
Two month postnatal			
Low SRQ score	-0.31(0.06)	-1.08(0.06)	-2.03(0.05)
High SRQ score	-0.94(0.25)	-1.15(0.25)	-2.15(0.24)
p-value*	0.02	0.79	0.63
Pregnancy or postnatal			
Low SRQ at all time point	-0.34(0.06)	-1.10(0.06)	-2.05(0.05)
High SRQ score at both time points	-0.69(0.23)	-1.17(0.15)	-2.10(0.19)
High SRQ score at Postnatal only	-1.24(0.49)	-1.11(0.41)	-2.25(0.27)
High SRQ score at Pregnancy only	-0.05(0.18)	-1.18(0.31)	-2.07(0.38)
p-value*	0.04	0.92	0.92

Table 4.9: Correlation of selected background parental characteristics and infants length-for-age z at two, six and 12 months of age

	Length-for-age					
Background characteristics	**Two months**		**Six months**		**Twelve months**	
	Correlation	**p-value**	**Correlation**	**p-value**	**Correlation**	**p-value**
Maternal age	-0.020	0.553	-0.004	0.910	-0.064	0.053
Age of father	0.016	0.648	-0.002	0.963	-0.043	0.198
Maternal MUAC	-0.023	0.498	0.109	0.001	0.079	0.016
Poverty index	-0.036	0.294	0.027	0.437	-0.127	0.000
Sanitary index	0.042	0.219	-0.038	0.266	-0.167	0.000
Maternal support	-0.039	0.246	-0.012	0.723	-0.053	0.111
Maternal autonomy	-0.022	0.507	-0.045	0.189	0.050	0.131

Bivariate associations of length-for-age z at two, six and twelve months of age and background characteristics of households and infants are summarized in table 4.10 Infant gender was significantly associated with mean length-for-age z at all time points female infants performing better at two, six and twelve months of age. Low birth weight and unavailability of toilet facility were associated with reduced mean value of length-for-age z at six and twelve months and rural residence was associated with reduced mean length-for-age z at two and twelve months of age. Having a family that uses water from an unprotected source, uses open field for rubbish disposal and having a non-literate father were significantly associated with reduced mean length-for-age z at twelve months of age. At two month follow-up the direction of association of mean length-for-age with the denial of colostrums and families' main source of water were counter intuitive.

Table 4.10: Mean(SD) of length-for-age z score of infants at two, six and 12 months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		*12 month follow-up*	
	Mean(SD)	p-value	Mean(SD)	p-value	Mean(SD)	p-value
Sex						
Male	-0.50(1.65)	0.004	-1.34(1.60)	0.000	-2.32(1.61)	0.000
Female	-0.18(1.64)		-0.81(1.50)		-1.74(1.39)	
Birth weight						
Normal	-0.37(1.65)	0.099	-0.90(1.51)	0.000	-1.90(1.52)	0.001
Low	-0.79(1.51)		-1.45(1.51)		-2.59(1.64)	
Missing	-0.22(1.67)		-1.36(1.66)		-2.21(1.51)	
Residence						
Rural	-0.41(1.67)	0.003	-1.11(1.56)	0.286	-2.14(1.47)	0.000
Urban	0.06(1.45)		-0.94(1.69)		-1.39(1.75)	
Under five children						
None	-0.27(1.53)	0.169	-1.07(1.63)	0.991	-1.85(1.64)	0.184
Only one	-0.38(1.64)		-1.08(1.61)		-2.09(1.53)	
Two or more	-0.33(1.75)		-1.09(1.47)		-2.08(1.45)	
Sanitation facility						
Toilet available	-0.34(1.62)	0.997	-0.96(1.54)	0.003	-1.91(1.57)	0.001
Open field	-0.34(1.70)		-1.29(1.62)		-2.25(1.44)	
Rubbish disposal						
Burn	-0.15(1.58)	0.060	-1.01(1.64)	0.477	-1.73(1.62)	0.001
Open field	-0.40(1.67)		-1.10(1.56)		-2.13(1.49)	
Water source						
Protected	-0.49(1.68)	0.000	-1.14(1.61)	0.151	-1.91(1.53)	0.000
unprotected	0.004(1.51)		-0.97(1.49)		-2.33(1.51)	
Initiation of breast feeding						
<= 1 hour	-0.28(1.64)	0.484	-1.02(1.61)	0.420	-2.08(1.62)	0.608
>1 hour	-0.37(1.66)		-1.12(1.56)		-2.02(1.49)	
Colostrums						
Given	-0.41(1.63)	0.008	-1.11(1.57)	0.361	-2.06(1.53)	0.463
Denied	-0.02(1.70)		-0.98(1.65)		-1.96(1.53)	
Pre-lacteal feeding						
Given	-0.80(1.73)	0.084	-1.13(1.44)	0.861	-2.23(1.42)	0.395
Not given	-0.32(1.65)		-1.09(1.59)		-2.03(1.53)	
Breastfeeding at two months						
Exclusive	-0.30(1.66)	0.069	-1.10(1.56)	0.597	-2.05(1.50)	0.695
Non-exclusive	-0.58(1.57)		-1.03(1.62)		-2.00(1.65)	
Maternal education						
Literate	-0.24(1.50)	0.116	-1.13(1.92)	0.670	-1.84(1.79)	0.053
Non-literate	-0.37(1.64)		-1.07(1.48)		-2.09(1.46)	
Father's education						
Literate	-0.31(1.67)	0.424	-1.06(1.58)	0.437	-1.96(1.56)	0.036
Non-literate	-0.41(1.61)		-1.15(1.58)		-2.19(1.46)	

SD: Standard deviation

Weight-for-age

Antenatal or postnatal CMD were not significantly associated with mean weight-for-age z of infants at the ages of two, six and twelve months of age (table 4.11). An increase in maternal MUAC and lower score on poor sanitary scale were significantly correlated with increased mean weight-for-age z score of infants at two, six and twelve months of age (table 4.12). Increased parental age and higher score on poverty index were inversely and significantly associated with increased mean weight-for-age z.

Table 4.11: Bivariate associations of prenatal maternal CMD and infant weight-for-age z score at two, six and 12 months of age

Timing of CMD measure and its categories	*2 months Mean(SE)*	*6 months Mean(SE)*	*12 months Mean(SE)*
Pregnancy			
Low SRQ score	-0.56(0.04)	-1.08(0.05)	-1.05(0.04)
High SRQ score	-0.41(0.12)	-1.20(0.14)	-1.16(0.14)
P-value*	0.19	0.38	0.37
Two month postnatal			
Low SRQ score	-0.56(0.04)	-1.10(0.04)	-1.07(0.04)
High SRQ score	-0.24(0.12)	-0.84(0.18)	-0.93(0.15)
p-value*	0.08	0.21	0.50
Pregnancy or postnatal			
Low SRQ at all time point	-0.57(0.04)	-1.11(0.05)	-1.06(0.05)
High SRQ score at both time points	-0.15(0.19)	-1.28(0.16)	-1.32(0.16)
High SRQ score at Postnatal only	-0.36(0.14)	-0.70(0.20)	-1.27(0.19)
High SRQ score at Pregnancy only	-0.48(0.14)	-0.93(0.27)	-0.67(0.20)
p-value*	0.27	0.95	0.12

SE: Standard error

Table 4.12: Correlation of selected background parental characteristics and infant weight-for-age at two, six and 12 months of age

	Weight-for-age					
Background	**Two months**		**Six months**		**Twelve months**	
characteristics	**Correlation**	**p-value**	**Correlation**	**p-value**	**Correlation**	**p-value**
Maternal age	-0.022	0.502	-0.073	0.030	-0.156	0.000
Age of father	-0.016	0.631	-0.100	0.003	-0.116	0.000
Maternal MUAC	0.084	0.012	0.164	0.000	0.113	0.000
Poverty index	-0.040	0.232	-0.163	0.000	-0.126	0.000
Sanitary index	-0.090	0.007	-0.229	0.000	-0.178	0.000
Maternal support	-0.022	0.515	-0.004	0.907	-0.021	0.534
Maternal autonomy	-0.022	0.501	0.033	0.331	0.016	0.634

Table 4.13: Mean(SD) of weight-for-age z score of infants at two, six and twelve months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		*12 month follow-up*	
	Mean(SD)	p-value	Mean(SD)	p-value	Mean(SD)	p-value
Sex						
Male	-0.67(1.16)	0.001	-1.26(1.37	0.000	-1.24(1.31)	0.000
Female	-0.41(1.10)		-0.91(1.12)		-0.88(1.26)	
Birth weight						
Normal	-0.47(1.14)	0.000	-1.05(1.24)	0.000	-0.98(1.30)	0.000
Low	-1.64(1.22)		-2.08(1.29)		-1.88(1.34)	
Missing	-0.53(1.05)		-1.01(1.25)		-1.10(1.24)	
Residence						
Rural	-0.56(1.14)	0.414	-1.19(1.23)	0.000	-1.12(1.31)	0.000
Urban	-0.47(1.13)		-0.51(1.31)		-0.66(1.12)	
Under five children						
None	-0.58(1.15)	0.641	-1.04(1.33)	0.821	-0.87(1.20)	0.075
Only one	-0.56(1.11)		-1.09(1.25)		-1.12(1.33)	
Two or more	-0.49(1.18)		-1.12(1.24)		-1.09(1.29)	
Sanitation facility						
Toilet available	-0.50(1.12)	0.120	-0.92(1.19)	0.000	-0.97(1.29)	0.006
Open field	-0.62(1.17)		-1.37(1.34)		-1.21(1.28)	
Rubbish disposal						
Burn	-0.43(1.12	0.099	-0.71(1.19)	0.000	-0.87(1.16)	0.021
Open field	-0.58(1.14)		-1.20(1.27)		-1.11(1.33)	
Water source						
Protected	-0.49(1.15)	0.022	-0.97(1.25)	0.000	-0.89(1.20)	0.000
Unprotected	-0.68(1.10)		-1.36(1.26)		-1.45(1.43)	
Initiation of breast feeding						
<= 1 hour	-0.41(1.12)	0.018	-0.89(1.23)	0.001	-0.94(1.36)	0.054
>1 hour	-0.60(1.14)		-1.18(1.23)		-1.12(1.26)	
Colostrums						
Given	-0.54(1.14)	0.697	-1.05(1.26)	0.062	-1.02(1.30)	0.050
Denied	-0.58(1.13)		-1.26(1.29)		-1.24(1.26)	
Pre-lacteal feeding						
Given	-0.61(0.97)	0.741	-1.23(1.41)	0.498	-1.15(1.18)	0.653
Not given	-0.54(1.15)		-1.09(1.26)		-1.06(1.30)	
Breastfeeding at two months						
Exclusive	-0.51(1.11)	0.034	-1.11(1.27)	0.228	-1.07(1.32)	0.500
Non-exclusive	-0.73(1.27)		-0.97(1.26)		-0.99(1.14)	
Maternal education						
Literate	-0.42(1.12)	0.116	-0.87(1.44)	0.012	-0.91(1.28)	0.076
Non-literate	-0.57(1.14)		-1.14(1.21)		-1.10(1.30)	
Father's education						
Literate	-0.52(1.15)	0.389	-1.00(1.25)	0.003	-0.99(1.30)	0.016
Non-literate	-0.59(1.10)		-1.28(1.29)		-1.21(1.28)	

SD: Standard deviation

Infants who are male, born with low birth weight, initiated breastfeeding later than an hour, and who had families that uses water from unprotected source had significantly smaller mean weight-for-age z score at 2, 6 and 12 months of age. Similarly, rural residence, not having a toilet facility, using an open field for rubbish disposal and having

non-literate father were significantly associated with lower infant mean weight-for-age z score at 6 and 12 months of age. Non-exclusive breastfeeding before two months of age was significantly associated with reduced infant weight-for-age z at the age of two months**.**

Stunting and underweight

Bivariate associations of infants' nutritional status (stunting and underweight) and level of CMD are presented in tables 4.14 and 4.15. CMD was not significantly associated with infant underweight and stunting at either two, six or twelve months of age, whether the level of CMD was measured during pregnancy, at two months postnatally, or according to the course of CMD across these two time points.

Table 4.14: Bivariate associations of maternal CMD and stunting of infants at two, six and 12 months of age

Timing of CMD measure and its categories	*2 months N(%)*	*6 months N(%)*	*12 months N(%)*
Pregnancy			
Low SRQ score	113(14.7%)	205(26.9)	386(47.8)
High SRQ score	14(13.6%)	25(25.0)	55(50.9)
P-value*	0.88	0.72	0.54
Two month postnatal			
Low SRQ score	120(14.5%)	220(26.8)	421(48.3)
High SRQ score	7(17.5%)	10(25.6)	18(43.9)
p-value*	0.65	1.00	0.63
Pregnancy or postnatal			
Low SRQ at all time point	108(14.4%)	201(26.9)	375(47.7)
High SRQ score at both time points	2(9.1%)	6(25.0)	9(39.1)
High SRQ score at Postnatal only	5(27.8%)	4(26.7)	9(50.0)
High SRQ score at Pregnancy only	12(14.8%)	19(25.0)	46(54.1)
p-value*	0.40	0.99	0.56

Table 4.15: Bivariate associations of maternal CMD and infant underweight at 2, 6 and 12 months of age

Timing of CMD measure and its categories	*2 months N(%)*	*6 months N(%)*	*12 months N(%)*
Pregnancy			
Low SRQ score	89(11.2%)	161(21.0)	167(20.5)
High SRQ score	8(7.6%)	27(27.6)	28(25.7)
P-value*	0.32	0.15	0.21
Two month postnatal			
Low SRQ score	96(11.2%)	182(22.0)	188(21.4)
High SRQ score	1(2.4%)	6(15.8)	6(14.3)
p-value*	0.12	0.43	0.34
Pregnancy or postnatal			
Low SRQ at all time point	89(11.5%)	160(21.3)	162(20.5)
High SRQ score at both time points	1(4.4%)	5(21.7)	2(8.3)
High SRQ score at Postnatal only	0(0.0%)	1(6.7)	4(22.2)
High SRQ score at Pregnancy only	7(8.5%)	22(29.3)	26(30.6)
p-value*	0.32	0.22	0.07

Bivariate associations of infant nutritional status (stunting and underweight) and selected background characteristics are summarized in tables 4.16-4.18. Male infants compared to females and infants born with low birth weight compared to infants born with normal birth weight had significantly increased prevalence of stunting and underweight, respectively, at two, six and twelve months of age. Prevalence of stunting at six and twelve months follow-up was also significantly higher among low birth weight infants, if the family uses unprotected water or do not have toilet facility. Similarly, prevalence of underweight was significantly higher among males than females at two and six months of age.

Significant association between prevalence of stunting at two months of follow-up and use of unprotected water by the family was counter intuitive. Prevalence of infant underweight was significantly higher at six and twelve month follow-up in the rural areas and if parents use unprotected water. Rural residence significantly increased the risk of stunting at two and twelve months of age but not at six months of age. Not having a toilet facility has significantly increased the risk of infant underweight at six months of age. Disposing of rubbish on the field is significantly associated with prevalence of underweight at 6 month and with that of stunting at 12 month. Delayed initiation of

breast feeding, denial of colostrums and having non-literate father were significantly associated with the increased prevalence of underweight at six months

Table 4.16: Prevalence of stunting of infants at two, six and twelve months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		*12 month follow-up*	
	Number (%)	p-value	Number (%)	p-value	Number (%)	p-value
Sex						
Male	77(17.5)	0.013	150(33.9)	0.000	266(57.0)	0.000
Female	50(11.6)		80(19.1)		175(39.0)	
Birth weight						
Normal	86(15.8)	0.356	121(23.0)	0.000	121(23.0)	0.008
Low	6(15.8)		14(31.8)		14(31.8)	
Missing	35(12.1)		95(32.7)		95(32.7)	
Residence						
Rural	119(16.0)	0.004	196(26.5)	0.795	398(50.2)	0.002
Urban	8(6.2)		32(25.6)		43(35.0)	
Under five children						
None	23(12.8)	0.749	48(27.1)	0.873	84(44.9)	0.590
Only one	67(15.1)		121(27.2)		230(49.4)	
Two or more	37(14.9)		61(25.4)		127(48.3)	
Sanitation facility						
Toilet available	80(14.4)	0.837	133(24.4)	0.047	257(44.6)	0.005
Open field	47(14.9)		97(30.6)		184(54.1)	
Rubbish disposal						
Burn	26(13.0)	0.475	53(27.6)	0.752	76(38.2)	0.001
Open field	101(15.0)		177(26.5)		365(51.0)	
Water source						
Protected	107(17.5)	0.000	175(29.5)	0.006	285(44.7)	0.001
Unprotected	20(7.7)		55(20.6)		156(56.5)	
Initiation of breast feeding						
<= 1 hour	37(13.5)	0.502	69(25.2)	0.439	142(49.5)	0.655
>1 hour	89(15.2)		159(27.7)		294(47.9)	
Colostrums						
Given	104(14.8)	0.768	189(27.4)	0.505	363(49.2)	0.313
Denied	22(13.8)		40(24.8)		74(44.9)	
Pre-lacteal feeding						
Given	8(21.6)	0.220	10(25.0)	0.766	21(48.8)	0.946
Not given	118(14.3)		219(27.1)		414(48.3)	
Breastfeeding at two months						
Exclusive	106(14.5)	0.888	192(26.8)	0.916	361(47.1)	0.137
Non-exclusive	21(15.0)		38(26.4)		78(53.8)	
Maternal education						
Literate	21(11.9)	0.271	52(30.2)	0.239	79(43.7)	0.176
Non-literate	106(15.2)		178(25.8)		362(49.3)	
Father's education						
Literate	21(11.9)	0.271	151(25.8)	0.378	292(46.7)	0.183
Non-literate	106(15.2)		78(28.7)		146(51.4)	

Table 4.17: Prevalence of underweight of infants at two, six and twelve months of age stratified by selected maternal, infant and environmental characteristics

Selected background characteristics	*2 month follow-up*		*6 month follow-up*		*12 month follow-up*	
	Number (%)	p-value	Number (%)	p-value	Number (%)	p-value
Sex						
Male	62(13.7)	0.005	119(26.9)	0.000	107(23.0)	0.173
Female	35(7.9)		69(16.4)		88(19.3)	
Birth weight						
Normal	57(10.2)	0.000	108(20.4)	0.000	108(20.4)	0.000
Low	14(32.6)		23(51.1)		23(51.1)	
Missing	26(8.8)		57(19.7)		57(19.7)	
Residence						
Rural	81(10.3)	0.527	173(23.4)	0.005	181(22.7)	0.002
Urban	16(12.4)		15(12.1)		14(11.3)	
Under five children						
None	22(11.9)	0.848	44(25.1)	0.429	28(14.8)	0.057
Only one	49(10.7)		91(20.4)		107(22.8)	
Two or more	26(10.2)		58(21.8)		60(22.8)	
Sanitation facility						
Toilet available	56(9.8)	0.205	94(17.2)	0.000	112(19.4)	0.088
Open field	41(12.5)		94(29.7)		83(24.1)	
Rubbish disposal						
Burn	25(12.4)	0.417	29(15.1)	0.011	34(17.1)	0.111
Open field	72(10.4)		159(23.7)		161(22.3)	
Water source						
Protected	65(10.4)	0.550	117(19.6)	0.020	104(16.2)	0.000
Unprotected	32(11.8)		71(26.7)		91(32.6)	
Initiation of breast feeding						
<= 1 hour	27(9.6)	0.515	47(17.1)	0.023	55(19.0)	0.281
>1 hour	67(11.1)		138(24.0)		137(22.2)	
Colostrums						
Given	77(10.7)	0.932	142(20.5)	0.040	153(20.6)	0.318
Denied	18(10.9)		45(28.0)		40(24.1)	
Pre-lacteal feeding						
Given	4(10.8)	0.992	9(22.5)	0.934	8(19.1)	0.728
Not given	91(10.8)		178(22.0)		184(21.3)	
Breastfeeding at two months						
Exclusive	76(10.1)	0.116	158(21.9)	0.835	167(21.6)	0.394
Non-exclusive	21(14.6)		30(21.1)		27(18.5)	
Maternal education						
Literate	21(10.9)	0.590	38(22.0)	0.934	36(19.5)	0.529
Non-literate	76(10.5)		150(21.7)		159(21.6)	
Father's education						
Literate	21(11.9)	0.590	117(19.8)	0.044	128(20.3)	0.407
Non-literate	76(10.5)		70(25.9)		65(22.7)	

There was no significant difference in the mean age of parents, maternal characteristics (i.e. MUAC, degree of autonomy, availability of support, and household characteristics (poor sanitary condition scale and poverty index) between families whose infants were undernourished at two months of age and those who were not (Table 4.18). There was a significant difference in the mean score of maternal MUAC and poor sanitary condition

scale between infants who were underweight at six and at twelve months follow-up compared to infants who were not underweight at these time points. Similarly, there was a significant difference in mean age of parents whose infants were underweight at one year follow-up and in the mean poverty score of households whose infants were underweight at six month follow-up compared to those whose infants were not underweight on the corresponding follow-up time points. Stunting of infants was only significantly associated with increased mean poverty score and increased mean of poor sanitary condition scale at one year follow-up.

Table 4.18: Mean (standard deviation) of selected background parental characteristics stratified by infants' nutritional status at 2, 6 and 12 months of age

Two month time point						
Background	**Underweight**			**Stunted**		
characteristics	**Yes**	**No**	**p-value**	**Yes**	**No**	**p-value**
Maternal age	26.4(6.2)	26.9(6.2)	0.488	26.7(6.3)	27.0(6.2)	0.634
Age of father	35.3(8.4)	36.1(8.7)	0.435	35.8(8.2)	36.1(8.8)	0.707
Maternal MUAC	24.5(2.0)	24.7(2.1)	0.429	24.6(2.0)	24.7(2.1)	0.729
Poverty index	7.1(1.9)	7.3(2.0)	0.479	7.5(1.8)	7.2(2.1)	0.102
Sanitary index	1.5(1.0)	1.4(0.9)	0.559	1.3(0.9)	1.4(1.0)	0.166
Maternal support	1.1(1.0)	1.1(1.1)	0.769	1.2(1.0)	1.1(1.1)	0.698
Maternal autonomy	1.0(1.9)	1.1(1.8)	0.504	1.1(1.8)	1.1(1.8)	0.924
Six month time point						
Maternal age	27.2(6.3)	26.7(6.2)	0.336	26.8(6.2)	26.9(6.2)	0.896
Age of father	37.1(9.2)	35.8(8.7)	0.071	36.0(8.8)	36.0(8.9)	0.994
Maternal MUAC	24.3(1.9)	24.8(2.2)	0.002	24.5(2.0)	24.8(2.3)	0.131
Poverty index	7.6(1.8)	7.2(2.0)	0.013	7.1(2.1)	7.3(2.0)	0.261
Sanitary index	1.7(1.0)	1.4(1.0)	0.000	1.4(1.0)	1.5(1.0)	0.665
Maternal support	1.10(0.97)	1.09(1.09)	0.866	1.09(1.05)	1.10(1.07)	0.967
Maternal autonomy	1.17(1.86)	1.14(1.87)	0.785	1.32(1.98)	1.08(1.83)	0.094
Twelve month time point						
Maternal age	28.0(6.4)	26.6(6.2)	0.004	27.2(6.1)	26.6(6.4)	0.165
Age of father	37.4(9.7)	35.7(8.6)	0.016	36.4(8.8)	35.8(8.9)	0.333
Maternal MUAC	24.4(2.0)	24.8(2.2)	0.013	24.6(2.1)	24.8(2.2)	0.105
Poverty index	7.5(1.7)	7.2(2.0)	0.078	7.4(1.8)	7.1(2.1)	0.022
Sanitary index	1.7(1.0)	1.4(0.9)	0.000	1.6(0.9)	1.3(0.9)	0.000
Maternal support	1.17(1.03)	1.07(1.08)	0.254	1.13(1.06)	1.05(1.07)	0.242
Maternal autonomy	1.18(1.88)	1.11(1.83)	0.650	1.06(1.82)	1.15(1.83)	0.444

4.3: Unadjusted and adjusted effects of household characteristics, maternal characteristic, infant characteristics and early feeding practices of the mother on infant growth.

Length

Crude and adjusted regression coefficients and their 95% confidence intervals for factors associated with infant length are summarized in tables 4.19-4.21. Female gender, low birth weight, rural residence and higher score on poor sanitary condition scale were significantly associated with shorter infant length at 2, 6 and 12 months of age. The exception was a positive association between higher score on poor sanitary condition scale and increased infant length at two month follow-up which was significant in the adjusted model. Not having obstetric complication during delivery only in bivariate analyses, increased maternal MUAC before and after adjusting for other risk factors and having younger mother after adjusting for other factors were significantly associated with increased infant length at two, six , and twelve months of age, respectively. Contrary to our expectation having a less autonomous mother was significantly associated with increased infant length independent of other factors at six months of age. In the bivariate analysis but not in the multivariable analysis increase in the poverty index was significantly associated with shorter infant length at two and twelve months of age. At six months of age increased score on the poverty index was significantly associated with increased length in the adjusted model but there was no significant association in the unadjusted analysis.

Weight

Results of bivariate and multivariable regression analysis which assessed the effects of several risk factors on infant weight measured at 2, 6 and 12 months of age are summarized in tables 4.19-4.21. Low birth weight, female gender, absence of obstetric complication during delivery, and reduced maternal MUAC were significantly associated with light infant weight at two, six and twelve month follow-up before and after adjusting for other risk factors. Similarly, increased maternal age and higher score on poor sanitary condition scale were significantly associated with light infant weight at six and twelve months of age. In bivariate analysis denial of colostrums, rural residence,

increased age of the father, scoring higher on poverty index were significantly associated with reduced infant weight at six and twelve months of age. In bivariate analysis but not in fully adjusted models there was a significant inverse association between delayed initiation of breast feeding and infant weight at two months, a positive association between maternal height and infant weight at two months, and a negative association between having more than two under five siblings and infant weight at six months of age.

Table 4.19: Predictors of infant growth at two months of age in Butajira Birth cohort, Ethiopia

Characteristics considered	Weight			Length		
	n (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)	N (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	146(15.9)	-0.02(-0.17,0.14)	-0.02(-0.17,0.14)	145(15.9)	-0.25(-1.03,0.52)	-0.60(-1.41,0.20)
No pre-lacteal food	864(95.4)	0.03(-0.25,0.32)	0.01(-0.28,0.30)	857(95.2)	0.18(-1.22,1.58)	0.19(-1.29,1.68)
Colostrums not given	166(18.3)	-0.13(-0.28,0.01)	-0.04(-0.19,0.11)	165(18.3)	0.19(-0.54,0.92)	0.53(-0.25,1.31)
Breast feeding delayed for 1 hour	618(68.1)	-0.15(-0.28,-0.03)	-0.12(-0.24,0.01)	614(68.2)	-0.51(-1.12,0.09)	-0.46(-1.11,0.18)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.9(6.2)	0.003(-0.01,0.01)	0.003(-0.01,0.02)	26.9(6.2)	0.01(-0.03,0.06)	-0.004(-0.07, 0.07)
Height (metres)	1.58(0.06)	0.01(0.002,0.02)	0.01(-0.00,0.02)	1.58(0.06)	0.01(-0.04,0.06)	0.01(-0.04,0.06)
Mid upper arm circumference (cm)	24.7(2.1)	0.05(0.02,0.08)	0.05(0.02,0.08)	24.7(2.1)	0.09(-0.05,0.22)	0.11(-0.04,0.25)
Being in polygamous marriage	162(17.6)	0.04(-0.11,0.18)	0.08(-0.09,0.25)	162(17.7)	-0.57(-1.31,0.16)	-0.77(-1.66,0.13)
Autonomy scale (0-5)	1.1(1.8)	-0.01(-0.04,0.02)	-0.02(-0.05,0.01)	1.1(1.8)	-0.09(-0.24,0.07)	-0.13(-0.03, 0.03)
Use khat and/or alcohol	104(11.3)	0.07(-0.11,0.25)	0.06(-0.12,0.24)	104(11.4)	-0.14(-1.03,0.75)	-0.02(-0.94,0.91)
Had at least one obstetric complication	573(64.5)	0.15(0.03,0.27)	0.14(0.02,0.26)	571(64.7)	-0.61(-1.21,-0.01)	-0.40(-1.03,0.23)
Household characteristics						
Urban residence	124(13.5)	0.03(-0.13,0.19)	-0.09(-0.33,0.15)	123(13.4)	1.20(0.40,2.00)	1.27(0.02,2.53)
Number of under 5 children: 0	189(20.5)	1	1	187(20.4)	1	1
1	470(51.0)	0.09(-0.06,0.24)	0.04(-0.12,0.20)	466(50.9)	0.02(-0.72,0.77)	-0.21(-1.04,0.62)
≥ 2	263(28.5)	0.06(-0.10,0.22)	-0.02(-0.19,0.16)	263(28.7)	-0.44(-1.26,0.38)	-0.66(-1.58,0.25)
Age of father in years	36.0(8.8)	-0.00(-0.01,0.01)	-0.004(-0.01,0.01)	36.1(8.9)	0.01(-0.03,0.004)	0.03(-0.03,0.08)
Poverty index (0 to 11)	7.3(2.0)	-0.02(-0.05,0.01)	-0.01(-0.05,0.03)	7.3(2.0)	-0.15(-0.29,-0.01)	-0.05(-0.26, 0.17)
Poor sanitary condition (0 to 3)	1.5(1.0)	-0.06(-0.12,0.002)	-0.04(-0.11,0.03)	1.5(0.9)	0.19(-0.11,0.49)	0.48(0.12, 0.84)
Higher level of social support (0 to 4)	1.1(1.1)	-0.02(-0.08,0.03)	-0.01(-0.06,0.04)	1.1(1.1)	0.06(-0.21,0.33)	0.13(-0.15, 0.40)
Characteristics of index child						
Female gender	456(49.5)	-0.32(-0.43,-0.21)	-0.31(-0.42,-0.19)	449(49.0)	-0.91(-1.47,-0.34)	-0.72(-1.32,-0.13)
Not immunised at two months	381(41.5)	-0.07(-0.19,0.05)	-0.10(-0.21,0.02)	379(41.6)	-0.26(-0.84,0.32)	-0.18(-0.78, 0.43)
Severe illness in the first 2 months	193(21.0)	-0.03(-0.17,0.11)	-0.04(-0.18,0.10)	192(21.1)	-0.21(-0.91,0.50)	-0.02(-0.76, 0.72)
Birth weight: Normal (>=2500gm)	570(61.8)	1	1	568(62.0)	1	1
Low (<2500gm)	43(4.7)	-0.63(-0.90,-0.37)	-0.67(-0.93,-0.40)	41(4.5)	-1.86(-3.21,-0.51)	-2.18(-3.61,-0.75)
Not measured	309(33.5)	0.06(-0.06,0.18)	0.08(-0.04,0.21)	307(35.5)	0.78(0.18,1.39)	0.78(0.13, 1.43)

Table 4.20: Predictors of infant growth at six months of age in Butajira Birth cohort, Ethiopia

Characteristics considered	Weight			Length		
	n (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)	N (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	146(15.9)	0.12(-0.11,0.34)	0.07(-0.15,0.30)	145(15.9)	0.25(-0.37,0.86)	0.39(-0.26,1.04)
No pre-lacteal food	864(95.4)	0.15(-0.23,0.53)	0.24(-0.14,0.61)	857(95.2)	0.07(-1.01,1.15)	0.08(-1.05,1.21)
Colostrums not given	166(18.3)	-0.21(-0.42,-0.001)	-0.08(-0.29,0.14)	165(18.3)	0.32(-0.27,0.90)	0.28(-0.34,0.90)
Breast feeding delayed for 1 hour	618(68.1)	-0.17(-0.34,0.01)	-0.12(-0.30,0.06)	614(68.2)	-0.30(-0.79,0.20)	0.45(-1.10,1.00)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.9(6.2)	-0.03(-0.04,-0.02)	-0.03(-0.06,-0.01)	26.9(6.2)	0.02(-0.02,0.06)	0.01(-0.04, 0.07)
Height (metres)	1.58(0.06)	0.01(-0.002,0.02)	0.01(-0.01,0.02)	1.58(0.06)	-0.01(-0.05,0.03)	-0.02(-0.06,0.02)
Mid upper arm circumference (cm)	24.7(2.1)	0.07(0.04,0.11)	0.07(0.03,0.11)	24.7(2.1)	0.18(0.07,0.29)	0.16(0.04,0.28)
Being in polygamous marriage	162(17.6)	-0.08(-0.29,0.14)	0.04(-0.21,0.29)	162(17.7)	0.04(-0.55,0.64)	0.01(-0.73,0.74)
Autonomy scale (0-5)	1.1(1.8)	0.01(-0.04,0.05)	-0.02(-0.06,0.03)	1.1(1.8)	-0.09(-0.21,0.04)	-0.14(-0.27, -0.01)
Use khat and/or alcohol	104(11.3)	0.18(-0.08,0.43)	0.11(-0.15,0.36)	104(11.4)	0.49(-0.22,1.21)	0.17(-0.58,0.93)
Had at least one obstetric complication	573(64.5)	0.22(0.05,0.39)	0.18(0.01,0.35)	571(64.7)	-0.02(-0.50,0.47)	-0.07(-0.57,0.44)
Household characteristics						
Urban residence	124(13.5)	0.40(0.17,0.64)	0.20(-0.15,0.55)	123(13.4)	0.49(-0.16, 1.15)	1.30(0.27,2.33)
Number of under 5 children: 0	189(20.5)	1	1	187(20.4)	1	1
1	470(51.0)	-0.15(-0.36,0.06)	0.07(-0.16,0.29)	466(50.9)	0.09(-0.51,0.69)	0.04(-0.63,0.72)
≥ 2	263(28.5)	-0.24(-0.48,-0.01)	-0.12(-0.37,0.13)	263(28.7)	0.09(-0.57,0.76)	-0.13(-0.88,0.61)
Age of father in years	36.0(8.8)	-0.02(-0.02,-01)	-0.003(-0.02,0.01)	36.1(8.9)	0.01(-0.01,0.04)	-0.01(-0.05,0.04)
Poverty index (0 to 11)	7.3(2.0)	-0.08(-0.12,-0.04)	0.003(-0.06,0.06)	7.3(2.0)	0.02(-0.10,0.13)	0.16(0.001, 0.35)
Poor sanitary condition (0 to 3)	1.5(1.0)	-0.20(-0.29,-0.12)	-0.12(-0.22,-0.02)	1.5(0.9)	-0.25(-0.48,-0.01)	-0.23(-0.52, 0.06)
Higher level of social support (0 to 4)	1.1(1.1)	-0.05(-0.13,0.03)	-0.04(-0.12,0.04)	1.1(1.1)	-0.06(-0.28,0.16)	-0.04(-0.26, 0.19)
Characteristics of index child						
Female gender	456(49.5)	-0.32(-0.48,-0.16)	-0.33(-0.49,-0.17)	449(49.0)	-0.74(-1.19,-0.28)	-0.66(-1.14,-0.18)
Not immunised at two months	381(41.5)	-0.08(-0.24,0.09)	-0.06(-0.22,0.11)	379(41.6)	0.30(-0.16,0.77)	0.31(-0.18, 0.80)
Severe illness in the first 2 months	193(21.0)	0.01(-0.19,0.21)	-0.05(-0.25,0.15)	192(21.1)	-0.20(-0.76,0.36)	-0.28(-0.87, 0.31)
Birth weight: Normal (>=2500gm)	570(61.8)	1	1	568(62.0)	1	1
Low (<2500gm)	43(4.7)	-0.77(-1.15,-0.38)	-0.73(-1.12,-0.34)	41(4.5)	-1.18(-2.22,-0.15)	-1.11(-2.21,-0.003)
Not measured	309(33.5)	-0.08(-0.25,0.09)	-0.12(-0.29,0.06)	307(35.5)	-0.86(-1.35,-0.37)	-0.92(-1.45, -0.40)

Table 4.21: Predictors of infant growth at 12 months of age in Butajira Birth cohort, Ethiopia

Characteristics considered	Weight			*Length*		
	n (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)	N (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	146(15.9)	0.12(-0.11,0.34)	0.07(-0.15,0.30)	145(15.9)	0.39(-0.28,1.06)	0.27(-0.42,0.96)
No pre-lacteal food	864(95.4)	0.15(-0.23,0.53)	0.24(-0.14,0.61)	857(95.2)	0.31(-0.84,1.45)	0.71(-0.47,1.89)
Colostrums not given	166(18.3)	-0.21(-0.42,-0.001)	-0.08(-0.29,0.14)	165(18.3)	0.26(-0.37,0.90)	0.39(-0.28,1.05)
Breast feeding delayed for 1 hour	618(68.1)	-0.17(-0.34,0.01)	-0.12(-0.30,0.06)	614(68.2)	0.18(-0.34,0.71)	0.45(-0.10,1.00)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.9(6.2)	-0.03(-0.04,-0.02)	-0.03(-0.05,-0.01)	26.9(6.2)	-0.03(-0.07,0.01)	-0.06(-0.12,- 0.003)
Height (metres)	1.58(0.06)	0.01(-0.002,0.02)	0.01(-0.01,0.02)	1.58(0.06)	0.02(-0.02,0.06)	-0.00(-0.04,0.04)
Mid upper arm circumference (cm)	24.7(2.1)	0.07(0.04,0.11)	0.07(0.03,0.11)	24.7(2.1)	0.12(0.01,0.23)	0.11(-0.01,0.24)
Being in polygamous marriage	162(17.6)	-0.08(-0.29,0.14)	0.04(-0.21,0.29)	162(17.7)	0.22(-0.42,0.86)	0.10(-0.67,0.87)
Autonomy scale (0-5)	1.1(1.8)	0.01(-0.04,0.05)	-0.02(-0.06,0.03)	1.1(1.8)	0.06(-0.07,0.20)	-0.02(-0.16, 0.12)
Use khat and/or alcohol	104(11.3)	0.18(-0.08,0.43)	0.11(-0.15,0.36)	104(11.4)	-0.10(-0.87,0.67)	-0.03(-0.82,0.76)
Had at least one obstetric complication	573(64.5)	0.22(0.05,0.39)	0.18(0.01,0.35)	571(64.7)	0.33(-0.18,0.85)	0.28(-0.25,0.82)
Household characteristics						
Urban residence	124(13.5)	0.40(0.17,0.64)	0.20(-0.15,0.55)	123(13.4)	1.59(0.88,2.30)	1.51(0.42,2.59)
Number of under 5 children: 0	189(20.5)	1	1	187(20.4)	1	1
1	470(51.0)	-0.15(-0.36,0.06)	0.07(-0.16,0.29)	466(50.9)	-0.49(-1.12,0.15)	-0.12(-0.84,0.59)
≥ 2	263(28.5)	-0.24(-0.48,-0.01)	-0.12(-0.37,0.13)	263(28.7)	-0.62(-1.33,0.09)	-0.48(-1.27,0.30)
Age of father in years	36.0(8.8)	-0.02(-0.02,-0.01)	-0.003(-0.02,0.01)	36.1(8.9)	-0.01(-0.04,0.02)	0.01(-0.03,0.06)
Poverty index (0 to 11)	7.3(2.0)	-0.08(-0.12,-0.04)	0.003(-0.06,0.06)	7.3(2.0)	-0.24(-0.36,-0.11)	0.02(-0.17, 0.20)
Poor sanitary condition (0 to 3)	1.5(1.0)	-0.21(-0.29,-0.12)	-0.12(-0.22,-0.02)	1.5(0.9)	-0.60(-0.85,-0.34)	-0.32(-0.63, -0.01)
Higher level of social support (0 to 4)	1.1(1.1)	-0.05(-0.13,0.03)	-0.04(-0.12,0.04)	1.1(1.1)	-0.22(-0.45,0.01)	-0.17(-0.41, 0.06)
Characteristics of index child						
Female gender	456(49.5)	-0.32(-0.48,-0.16)	-0.33(-0.49,-0.17)	449(49.0)	-0.77(-1.25,-0.28)	-0.76(-1.27,-0.26)
Not immunised at two months	381(41.5)	-0.08(-0.24,0.09)	-0.06(-0.22,0.11)	379(41.6)	0.19(-0.31,0.68)	0.15(-0.37, 0.66)
Severe illness in the first 2 months	193(21.0)	0.01(-0.19,0.21)	-0.05(-0.25,0.15)	192(21.1)	0.35(-0.25,0.95)	0.15(-0.47, 0.78)
Birth weight: Normal (>=2500gm)	570(61.8)	1	1	568(62.0)	1	1
Low (<2500gm)	43(4.7)	-0.77(-1.15,-0.38)	-0.73(-1.12,-0.34)	41(4.5)	-1.71(-2.89,-0.54)	-1.54(-2.77,-0.31)
Not measured	309(33.5)	-0.08(-0.25,0.09)	-0.12(-0.29,0.06)	307(35.5)	-0.80(-1.32,-0.28)	-1.03(-1.58, -0.48)

Length-for-age

Unadjusted and adjusted regression coefficients from models investigating predictors of length-for-age of infants at 2, 6 and 12 months of age are summarized in tables 4.22 – 4.24. Male gender and rural residence were significant predictors of reduced length-for-age at two, six and twelve month follow-up and reduced maternal MUAC and low birth weight were significantly associated with reduced length-for-for-age at six and twelve month follow-ups. Not receiving pre-lacteal food, denial of colostrums, and higher score on poor sanitary condition scale were significantly associated with an increase of length-for-age z score at two months of age. In bivariate analysis lower score on poverty scale and lower score on poor sanitary condition scale were significantly associated with an increase of length-for-age z score at twelve months of age and in the adjusted model association was only significant between poor sanitary condition scale and length-for-age z. Marginally significant inverse association in the adjusted model between maternal autonomy score and length-for-age z at six months of age was contrary to our expectation

Weight-for-age

Unadjusted and adjusted regression coefficients from models investigating predictors of weight-for-age z of infants at 2, 6 and 12 months of age are summarized in tables 4.22 – 4.24. Male gender, low birth weight, higher score on poor sanitary condition scale and lower maternal MUAC were significant predictors of lower weight-for-age z at two, six and twelve months of age. Increased maternal height, urban residence and at least one obstetric complication during delivery were significantly associated with increased weight-for-age z at 2, 6 and 12 months of age, respectively. Delayed invitation of breastfeeding was significant predictor of lower weight-for-age at 2 months of age in bivariate analysis and at 6 months of age before and after adjusting for other risk factors. In bivariate analysis reduced weight-for-age z was associated (a) at two months of age with non-exclusive breastfeeding and with mother not having at least obstetric complication during delivery, (b) at six months of age with having an aged father and with higher score on poverty index, (c) at one year of age with rural residence, having one under five sibling, having an aged father, scoring higher on poverty index, and denial of colostrums. These associations were not replicated in multivariable regression analysis. At one year of age increased maternal age was significantly associated with lower weight-for-age z before and after adjusting for confounding variables.

Table 4.22: Predictors of infant undernutrition at two months of age in Butajira Birth cohort, Ethiopia

	Weight-for-age			*Length-for-age*		
	N (%) or Mean(SD)	**Crude $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**	**N (%) or Mean(SD)**	**Crude $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	142 (16.5)	-0.22(-0.42,-0.02)	-0.15(-0.36,0.05)	144(16.8)	-0.28(-0.57, 0.02)	-0.29(-0.60,0.02)
No pre-lacteal food	810 (95.3)	0.06(-0.31, 0.44)	0.08(-0.30,0.46)	806(95.3)	0.48(-0.06, 1.03)	0.59(0.02,1.15)
Colostrums not given	160 (18.8)	-0.04(-0.23, 0.15)	0.10(-0.10,0.30)	160(18.9)	0.39(0.10, 0.67)	0.44(0.14,0.74)
Breast feeding delayed for 1 hour	576 (67.8)	-0.20(-0.46,-0.03)	-0.11(-0.27,0.06)	574(67.8)	-0.08(-0.32,0.15)	-0.02(-0.27,0.23)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.8(6.2)	-0.00(-0.01, 0.01)	-0.01(-0.02, 0.01)	26.9(6.2)	-0.00(-0.02,0.01)	-0.02(-0.05, 0.01)
Height (metres)	1.58(0.06)	0.02(0.00, 0.03)	0.02(0.01,0.03)	1.58(0.06)	0.01(-0.01,0.02)	0.01(-0.01,0.03)
Mid upper arm circumference (cm)	24.7(2.1)	0.05(0.02, 0.09)	0.05(0.02,0.09)	24.7(2.1)	-0.01(-0.07,0.04)	-0.01(-0.06,0.03)
Being in polygamous marriage	153 (17.9)	0.07(-0.12, 0.26)	0.08(-0.15,0.31)	153(17.7)	0.02(-0.06, 0.31)	-0.14(-0.48,0.20)
Autonomy scale (0-5)	1.1(1.9)	-0.02(-0.06, 0.02)	-0.03(-0.07,0.01)	1.1(1.9)	-0.02(-0.08,0.03)	-0.04(-0.10,0.03)
Use khat and/or alcohol	101(11.7)	0.05(-0.18,0.28)	0.02(-0.22,0.26)	100(11.6)	-0.31(-0.65, 0.03)	-0.26(-0.61,0.09)
Had at least one obstetric complication	534(63.9)	0.19(0.03, 0.34)	0.14(-0.02,0.30)	532(63.9)	-0.19(-0.42,0.04)	-0.13(-0.37,0.11)
Household characteristics						
Urban residence	124 (14.4)	0.09(-0.12, 0.30)	0.02(-0.30,0.34)	123(14.3)	0.47(0.16, 0.78)	0.76(0.28,1.24)
Number of under 5 children: 0	174 (20.1)	1	1	176(20.4)	1	1
1	447 (51.7)	0.02(-0.17, 0.21)	0.02(-0.19,0.25)	445(51.7)	-0.11(-0.40,0.18)	-0.07(-0.39,0.25)
≥ 2	243 (28.1)	0.09(-0.12, 0.31)	-0.02(-0.26,0.21)	240(27.9)	-0.06(-0.38,0.26)	-0.06(-0.41,0.29)
Age of father in years	36.1(8.8)	-0.00(-0.01, 0.01)	-0.00(-0.01,0.01)	36.0(8.9)	0.00(-0.01, 0.02)	0.02(-0.01,0.04)
Poverty index (0 - 11)	7.3(2.0)	-0.02(-0.06, 0.02)	0.01(-0.04,0.07)	7.3(2.0)	-0.03(-0.09,0.02)	0.02(-0.06,0.11)
Poor sanitary condition (0 -3)	1.5(1.0)	-0.11(-0.18, -0.03)	-0.07(-0.16,0.02)	1.5(1.0)	0.07(-0.05,0.18)	0.20(0.06,0.34)
Level of social support (0 - 4)	1.1(1.1)	-0.03(-0.10, 0.04)	-0.02(-0.09,0.05)	1.1(1.1)	-0.07(-0.18,0.03)	-0.06(-0.17,0.05)
Characteristics of index child						
Female gender	422 (48.8)	0.26(0.11, 0.41)	0.27(0.12,0.42)	420(48.8)	0.32(0.10,0.54)	0.38(0.15,0.60)
Not immunised at two months	352 (40.8)	-0.07(-0.22, 0.08)	-0.11(-0.26,0.05)	351(40.9)	-0.05(-0.27,0.17)	-0.02(-0.25,0.21)
Severe illness in the first 2 months	184 (21.4)	-0.16(-0.34, 0.03)	-0.11(-0.29,0.08)	184(21.4)	-0.11(-0.38,0.17)	0.04(-0.25,0.33)
Birth weight: Normal (>=2500g)	530 (61.3)	1	1	527(61.2)	1	1
Low (<2500g)	45 (5.2)	-1.17(-1.52, -0.83)	-1.21(-1.58,-0.85)	44(5.1)	-0.42(-0.96,0.12)	-0.48(-1.05,0.09)
Not measured	289 (33.5)	-0.06(-0.22, 0.10)	-0.04(-0.21,0.13)	290(33.7)	0.15(-0.08,0.39)	0.11(-0.14,0.36)

Table 4.23: Predictors of infant undernutrition at six months of age in Butajira Birth cohort, Ethiopia

	Weight-for-age			*Length-for-age*		
	N (%) or Mean(SD)	**Crude $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**	**N (%) or Mean(SD)**	**Crude $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	142 (16.5)	0.14(-0.09, 0.37)	0.13(-0.09,0.36)	144(16.8)	0.08(-0.21, 0.35)	0.20(-0.09, 0.50)
No pre-lacteal food	810 (95.3)	0.14(-0.26, 0.54)	0.21(-0.19,0.60)	806(95.3)	0.04(-0.46, 0.55)	0.04(-0.48, 0.56)
Colostrums not given	160 (18.8)	-0.21(-0.43, 0.01)	0.02(-0.20,0.24)	160(18.9)	0.13(-0.15, 0.40)	0.17(-0.11, 0.46)
Breast feeding delayed for 1 hour	576 (67.8)	-0.30(-0.48,-0.12)	-0.21(-0.39,-0.03)	574(67.8)	-0.09(-0.32,0.13)	-0.04(-0.28, 0.20)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.8(6.2)	-0.01(-0.03, 0.00)	-0.00(-0.02, 0.02)	26.9(6.2)	0.00(-0.02,0.02)	0.00(-0.03, 0.03)
Height (metres)	1.58(0.06)	0.00(-0.01, 0.02)	0.00(-0.01,0.02)	1.58(0.06)	-0.00(-0.02,0.02)	-0.01(-0.02, 0.01)
Mid upper arm circumference (cm)	24.7(2.1)	0.10(0.06, 0.14)	0.09(0.05,0.13)	24.7(2.1)	0.08(0.03, 0.13)	0.08(0.02, 0.13)
Being in polygamous marriage	153 (17.9)	0.04(-0.18, 0.26)	0.13(-0.13,0.38)	153(17.7)	0.11(-0.17, 0.38)	0.10(-0.23, 0.44)
Autonomy scale (0-5)	1.1(1.9)	0.02(-0.03, 0.06)	-0.02(-0.07,0.03)	1.1(1.9)	-0.05(-0.10, 0.01)	-0.07(-0.13,-0.01)
Use khat and/or alcohol	101(11.7)	0.14(-0.13,0.40)	0.10(-1.16,0.36)	100(11.6)	0.34(0.01, 0.67)	0.21(-0.13, 0.55)
Had at least one obstetric complication	534(63.9)	0.17(-0.01, 0.35)	0.11(-0.07,0.29)	532(63.9)	-0.00(-0.22,0.22)	-0.06(-0.29, 0.18)
Household characteristics						
Urban residence	124 (14.4)	0.67(0.44, 0.91)	0.44(0.08,0.80)	123(14.3)	0.16(-0.14, 0.46)	0.62(0.14, 1.09)
Number of under 5 children: 0	174 (20.1)	1	1	176(20.4)	1	1
1	447 (51.7)	-0.05(-0.27, 0.17)	0.13(-0.11,0.36)	445(51.7)	-0.01(-0.29,0.27)	0.05(-0.26, 0.36)
≥ 2	243 (28.1)	-0.08(-0.33, 0.17)	-0.08(-0.34,0.19)	240(27.9)	-0.02(-0.33,0.29)	-0.10(-0.44, 0.25)
Age of father in years	36.1(8.8)	-0.01(-0.02, -0.003)	-0.02(-0.03,0.00)	36.0(8.9)	0.00(-0.01, 0.01)	-0.01(-0.03, 0.01)
Poverty index (0 - 11)	7.3(2.0)	-0.10(-0.14, -0.06)	0.01(-0.05,0.07)	7.3(2.0)	0.02(-0.03, 0.07)	0.08(-0.00, 0.16)
Poor sanitary condition (0 -3)	1.5(1.0)	-0.30(-0.38, -0.21)	-0.20(-0.30,-0.10)	1.5(1.0)	-0.06(-0.17,0.05)	-0.05(-0.18, 0.09)
Level of social support (0 - 4)	1.1(1.1)	-0.00(-0.08, 0.08)	0.02(-0.06,0.10)	1.1(1.1)	-0.02(-0.12,0.08)	-0.01(-0.11, 0.09)
Characteristics of index child						
Female gender	422 (48.8)	0.36(0.19, 0.52)	0.37(0.20,0.54)	420(48.8)	0.53(0.32,0.74)	0.58(0.36, 0.80)
Not immunised at two months	352 (40.8)	0.03(-0.14, 0.21)	0.10(-0.07,0.27)	351(40.9)	0.09(-0.12, 0.31)	0.12(-0.11, 0.34)
Severe illness in the first 2 months	184 (21.4)	-0.12(-0.33, 0.08)	-0.01(-0.21,0.20)	184(21.4)	-0.15(-0.41,0.11)	-0.08(-0.35, 0.19)
Birth weight: Normal (>=2500g)	530 (61.3)	1	1	527(61.2)	1	1
Low (<2500g)	45 (5.2)	-1.03(-1.41, -0.65)	-0.97(-1.36,-0.59)	44(5.1)	-0.55(-1.03,-0.05)	-0.62(-1.13, -0.12)
Not measured	289 (33.5)	0.03(-0.14, 0.21)	0.01(-0.17,0.19)	290(33.7)	-0.45(-0.68,-0.23)	-0.45(-0.69,-0.21)

Table 4.24: Predictors of infant undernutrition at 12 months of age in Butajira Birth cohort, Ethiopia

Characteristics considered	Weight-for-age			*Length-for-age*		
	n (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)	N (%) or Mean(SD)	Crude $\hat{\beta}$ (95% CI)	Adjusted $\hat{\beta}$ (95% CI)
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	146(15.9)	-0.08(-0.15,0.31)	0.07(-0.16,0.30)	145(15.9)	0.05(-0.22,0.33)	0.02(-0.26, 0.29)
No pre-lacteal food	864(95.4)	0.09(-0.31,0.49)	0.18(-0.21,0.58)	857(95.2)	0.20(-0.26,0.67)	0.34(-0.12, 0.81)
Colostrums not given	166(18.3)	-0.22(-0.44,-0.00)	-0.08(-0.30,0.14)	165(18.3)	0.10(-0.16,0.36)	0.19(-0.08, 0.45)
Breast feeding delayed for 1 hour	618(68.1)	-0.18(-0.36,0.00)	-0.11(-0.29,0.08)	614(68.2)	0.06(-0.16,0.27)	0.17(-0.05, 0.39)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.9(6.2)	-0.03(-0.04,-0.02)	-0.03(-0.05, -0.01)	26.9(6.2)	-0.01(-0.03,0.00)	-0.02(-0.04, 0.00)
Height (metres)	1.58(0.06)	0.01(-0.01,0.02)	0.01(-0.01,0.02)	1.58(0.06)	0.01(-0.01,0.02)	0.00(-0.01,0.02)
Mid upper arm circumference (cm)	24.7(2.1)	0.07(0.03,0.11)	0.07(0.03,0.11)	24.7(2.1)	0.06(0.01,0.10)	0.06(0.01,0.10)
Being in polygamous marriage	162(17.6)	-0.02(-0.24,0.20)	0.05(-0.19,0.32)	162(17.7)	0.13(-0.13,0.38)	0.08(-0.22,0.39)
Autonomy scale (0-5)	1.1(1.8)	0.01(-0.04,0.05)	-0.01(-0.06, 0.03)	1.1(1.8)	0.04(-0.02,0.09)	0.01(-0.05, 0.06)
Use khat and/or alcohol	104(11.3)	0.21(-0.05,0.48)	0.12(-0.15,0.38)	104(11.4)	0.02(-0.30,0.33)	0.01(-0.31, 0.32)
Had at least one obstetric complication	573(64.5)	0.24(0.06,0.41)	0.18(0.01,0.36)	571(64.7)	0.07(-0.14,0.28)	0.04(-0.17, 0.26)
Household characteristics						
Urban residence	124(13.5)	0.46(0.22,0.71)	0.25(-0.12,0.61)	123(13.4)	0.75(0.46,1.03)	0.68(0.25,1.11)
Number of under 5 children: 0	189(20.5)	1	1	187(20.4)	1	1
1	470(51.0)	-0.25(-0.47,-0.03)	0.03(-0.20,0.27)	466(50.9)	-0.23(-0.49,0.03)	0.01(-0.28, 0.29)
≥ 2	263(28.5)	-0.22(-0.46,0.02)	-0.13(-0.39,0.13)	263(28.7)	-0.23(-0.51,0.06)	-0.16(-0.47, 0.15)
Age of father in years	36.0(8.8)	-0.02(-0.03,-0.01)	-0.00(-0.02,0.01)	36.1(8.9)	-0.01(-0.02,0.00)	0.00(-0.02, 0.02)
Poverty index (0 to 11)	7.3(2.0)	-0.08(-0.12,-0.04)	0.01(-0.05, 0.07)	7.3(2.0)	-0.10(-0.15,-0.05)	0.01(-0.06, 0.09)
Poor sanitary condition (0 to 3)	1.5(1.0)	-0.24(-0.33,-0.16)	-0.16(-0.26, -0.05)	1.5(0.9)	-0.26(-0.37,-0.16)	-0.13(-0.25,-0.01)
Higher level of social support (0 to 4)	1.1(1.1)	-0.04(-0.11,0.04)	-0.03(-0.10, 0.05)	1.1(1.1)	-0.08(-0.17,-0.02)	-0.02(-0.16, 0.03)
Characteristics of index child						
Female gender	456(49.5)	0.37(0.20,0.53)	0.35(0.18,0.52)	449(49.0)	0.58(0.38,0.77)	0.58(0.38, 0.78)
Not immunised at two months	381(41.5)	-0.09(-0.26,0.08)	-0.05(-0.22, 0.12)	379(41.6)	0.07(-0.13,0.27)	0.07(-0.14, 0.27)
Severe illness in the first 2 months	193(21.0)	-0.08(-0.29,0.13)	-0.07(-0.27, 0.14)	192(21.1)	0.01(-0.24,0.25)	0.06(-0.19,0.31)
Birth weight: Normal (>=2500gm)	570(61.8)	1	1	568(62.0)	1	1
Low (<2500gm)	43(4.7)	-0.90(-1.29,-0.50)	-0.85(-1.25,-0.45)	41(4.5)	-0.69(-1.17,-0.21)	-0.67(-1.16,-0.17)
Not measured	309(33.5)	-0.12(-0.29,0.06)	-0.12(-0.30, 0.06)	307(35.5)	-0.31(-0.52,-0.10)	-0.37(-0.59,-0.16)

Stunting and underweight

Results from bivariate analysis of each covariate and the growth outcomes (stunting and underweight) are presented in table 4.25-4.27. At two month follow-up male gender, low birth weight and being born without mother having obstetric complication were risk factors for underweight. Similarly, rural residence and male gender were risk factors for stunting at two months. Living rurally, having an older parent, scoring more on the sanitary scale and poverty index, being of male gender, low birth weight, not receiving colostrum, delayed initiation of breastfeeding, and lower maternal mid-upper arm circumference (MUAC) were significant predictors of underweight at the age of six months although the significance of MUAC was marginal. Only male gender and being vaccinated before two months of age were associated with stunting. At the age of twelve months, male gender, not being given colostrum, and delayed initiation of breastfeeding were no longer risk factors for being underweight, but having other children under the age of five became independently associated, along with older parental age, rural residence, poor sanitary conditions, low birth weight and maternal undernutrition. With respect to stunting at twelve months of age the risk factors were rural residence, increased values of poverty index and poor sanitary scores, male gender and low birth weight. There was very little difference in the magnitude and confidence intervals of the effect size of parental age on underweight and stunting at six and at twelve months of age

The adjusted effects of covariates on the infant growth outcomes (stunting and underweight) are presented in tables 4.25-4.27. At two month follow-up scoring high on poor sanitary condition scale and urban residence were protective against stunting, female gender was protective against underweight and stunting, and low birth weight was risk factor for underweight. Poor sanitation and low birth weight were associated with being underweight at both six and twelve months of age and with stunting at twelve months of age. Male gender was associated with stunting at both six and twelve months of age and being underweight at six months. Rural residence was associated with stunting and underweight at twelve months of age but not associated with infant undernutrition at six months of age. Infants with older parents were more likely to be disadvantaged at six and twelve months of age although the effect size associated with one year increase in parental age was relatively small. Maternal nutritional status was associated with being underweight at both six and twelve months of age but not with stunting. Having a sibling

aged under five and a higher score on the poverty scale were protective against becoming underweight at the six and twelve month time points, respectively, with marginal statistical significance of the latter. A higher value on the autonomy scale was associated with stunting at six month although the statistical significance of the effect was marginal.

Table 4.25: Predictors of infant undernutrition at two months of age in Butajira Birth cohort, Ethiopia

	Underweight			*Stunting*		
	N (%) or Mean(SD)	**Crude OR (95% CI)**	**Adjusted OR (95% CI)**	**N (%) or Mean(SD)**	**Crude OR (95% CI)**	**Adjusted OR (95% CI)**
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	142 (16.5)	1.51(0.90,2.55)	1.28(0.70,2.34)	144(16.8)	1.04(0.62,1.72)	1.15(0.66,2.01)
No pre-lacteal food	810 (95.3)	0.99(0.34,2.87)	1.26(0.72,2.22)	806(95.3)	0.61(0.27,1.36)	0.45(0.18,1.16)
Colostrums not given	160 (18.8)	1.02(0.59, 1.76)	0.80(0.43,1.50)	160(18.9)	0.93(0.57,1.52)	0.87(0.50,1.49)
Breast feeding delayed for 1 hour	576 (67.8)	1.17(0.73, 1.87)	1.14(0.67,1.93)	574(67.8)	1.15(0.76,1.74)	1.17(0.73,1.85)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.8(6.2)	0.99(0.95,1.02)	1.01(0.95,1.06)	26.9(6.2)	0.99(0.96,1.02)	1.00(0.95,1.05)
Height (metres)	1.58(0.06)	0.98(0.94,1.01)	0.98(0.94,1.02)	1.58(0.06)	0.99(0.96,1.02)	0.98(0.95,1.01)
Mid upper arm circumference (cm)	24.7(2.1)	0.96(0.87, 1.06)	0.92(0.82,1.04)	24.7(2.1)	0.98(0.90,1.08)	0.96(0.87,1.07)
Being in polygamous marriage	153 (17.9)	0.86(0.49, 1.52)	0.92(0.43,1.93)	153(17.7)	0.76(0.45,1.28)	0.88(0.46,1.69)
Autonomy scale (0-5)	1.1(1.9)	0.96(0.85,1.08)	0.96(0.84,1.10)	1.1(1.9)	0.99(0.90,1.10)	1.00(0.89,1.13)
Use khat and/or alcohol	101(11.7)	1.10(0.85,2.09)	1.24(0.61,2.54)	100(11.6)	1.40(0.82,2.40)	1.32(0.73,2.39)
Had at least one obstetric complication	534(63.9)	0.62(0.40,0.96)	0.67(0.41,1.09)	532(63.9)	1.05(0.70,1.56)	0.92(0.59,1.42)
Household characteristics						
Urban residence	124 (14.4)	1.20(0.68, 2.12)	1.18(0.43,3.21)	123(14.3)	0.35(0.17,0.73)	0.17(0.06,0.49)
Number of under 5 children: 0	174 (20.1)	1	1	176(20.4)	1	1
1	447 (51.7)	0.89(0.52, 1.52)	1.13(0.57,2.22)	445(51.7)	1.21(0.73,2.02)	1.09(0.60,1.99)
≥ 2	243 (28.1)	0.84(0.46, 1.54)	1.26(0.60,2.65)	240(27.9)	1.19(0.68,2.09)	0.90(0.46,1.74)
Age of father in years	36.1(8.8)	0.99(0.97,1.02)	0.99(0.95,1.04)	36.0(8.9)	1.00(0.97,1.02)	1.00(0.96,1.04)
Poverty index (0 - 11)	7.3(2.0)	0.96(0.87,1.07)	0.92(0.78,1.09)	7.3(2.0)	1.09(0.98,1.20)	1.00(0.85,1.16)
Poor sanitary condition (0 -3)	1.5(1.0)	1.07(0.86,1.33)	1.16(0.87,1.55)	1.5(1.0)	0.87(0.71,1.06)	0.64(0.50,0.83)
Level of social support (0 - 4)	1.1(1.1)	1.03(0.84,1.26)	1.00(0.80,1.24)	1.1(1.1)	1.04(0.87,1.24)	1.01(0.83,1.22)
Characteristics of index child						
Female gender	422 (48.8)	0.54(0.35,0.83)	0.50(0.31,0.81)	420(48.8)	0.62(0.42,0.90)	0.57(0.38,0.87)
Not immunised at two months	352 (40.8)	0.89(0.57,1.37)	1.06(0.65,1.71)	351(40.9)	1.06(0.72,1.55)	1.06(0.70,1.60)
Severe illness in the first 2 months	184 (21.4)	1.38(0.85,2.25)	1.23(0.70,2.15)	184(21.4)	1.13(0.71,1.78)	0.86(0.51,1.44)
Birth weight: Normal (>=2500g)	530 (61.3)	1	1	527(61.2)	1	1
Low (<2500g)	45 (5.2)	4.25(2.12,8.51)	5.20(2.41,11.20)	44(5.1)	1.00(0.41,2.47)	1.02(0.37,2.83)
Not measured	289 (33.5)	0.85(0.52,1.38)	0.72(0.41,1.25)	290(33.7)	0.74(0.48,1.12)	0.83(0.52,1.33)

Table 4.26: Predictors of infant undernutrition at six months of age in Butajira Birth cohort, Ethiopia

	Underweight			*Stunting*		
	N (%) or Mean(SD)	**Crude OR (95% CI)**	**Adjusted OR (95% CI)**	**N (%) or Mean(SD)**	**Crude OR (95% CI)**	**Adjusted OR (95% CI)**
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	142 (16.5)	0.95(0.61,1.48)	0.94(0.57,1.53)	144(16.8)	0.98(0.65,1.47)	0.80(0.51,1.24)
No pre-lacteal food	810 (95.3)	0.97(0.45,2.07)	1.08(0.44,2.63)	806(95.3)	1.12(0.54,2.32)	1.21(0.54,2.72)
Colostrums not given	160 (18.8)	1.50(1.02,2.22)	1.06(0.67,1.66)	160(18.9)	0.87(0.59,1.30)	0.85(0.55,1.31)
Breast feeding delayed for 1 hour	576 (67.8)	1.53(1.06,2.21)	1.40(0.93,2.12)	574(67.8)	1.14(0.82,1.58)	1.13(0.79,1.63)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.8(6.2)	1.01(0.99,1.04)	1.01(0.97,1.05)	26.9(6.2)	1.00(0.97,1.02)	0.99(0.96,1.03)
Height (metres)	1.58(0.06)	0.99(0.97,1.02)	1.00(0.97,1.03)	1.58(0.06)	1.01(0.98,1.03)	1.00(0.98,1.03)
Mid upper arm circumference (cm)	24.7(2.1)	0.88(0.81,0.96)	0.87(0.80,0.96)	24.7(2.1)	0.95(0.88,1.02)	0.94(0.87,1.02)
Being in polygamous marriage	153 (17.9)	1.01(0.67,1.54)	0.89(0.52,1.54)	153(17.7)	0.98(0.66,1.45)	1.06(0.64,1.76)
Autonomy scale (0-5)	1.1(1.9)	1.01(0.93,1.10)	1.06(0.96,1.17)	1.1(1.9)	1.07(0.99,1.16)	1.11(1.02,1.21)
Use khat and/or alcohol	101(11.7)	0.76(0.44,1.29)	0.84(0.47,1.52)	100(11.6)	0.90(0.56,1.46)	1.15(0.69,1.93)
Had at least one obstetric complication	534(63.9)	0.82(0.59,1.15)	0.97(0.66,1.42)	532(63.9)	0.82(0.60,1.12)	0.86(0.61,1.21)
Household characteristics						
Urban residence	124 (14.4)	0.45(0.26, 0.80)	0.54(0.23,1.27)	123(14.3)	1.06(0.69,1.62)	0.54(0.27,1.09)
Number of under 5 children: 0	174 (20.1)	1	1	176(20.4)	1	1
1	447 (51.7)	0.76(0.50, 1.15)	0.54(0.33,0.89)	445(51.7)	1.00(0.68,1.48)	0.95(0.60,1.51)
≥ 2	243 (28.1)	0.83(0.53, 1.31)	0.73(0.43,1.27)	240(27.9)	0.92(0.59,1.42)	0.99(0.59,1.66)
Age of father in years	36.1(8.8)	1.02(1.00,1.04)	1.03(1.00,1.06)	36.0(8.9)	1.00(0.98,1.02)	1.01(0.98,1.04)
Poverty index (0 - 11)	7.3(2.0)	1.02(1.02,1.22)	0.97(0.85,1.11)	7.3(2.0)	0.96(0.89,1.03)	0.94(0.83,1.06)
Poor sanitary condition (0 -3)	1.5(1.0)	1.46(1.23,1.74)	1.35(1.09,1.68)	1.5(1.0)	0.97(0.83,1.13)	0.96(0.79,1.16)
Level of social support (0 - 4)	1.1(1.1)	1.01(0.87,1.18)	1.00(0.84,1.18)	1.1(1.1)	1.00(0.86,1.15)	1.00(0.86,1.16)
Characteristics of index child						
Female gender	422 (48.8)	0.53(0.38,0.74)	0.49(0.34,0.71)	420(48.8)	0.46(0.33,0.63)	0.40(0.28,0.56)
Not immunised at two months	352 (40.8)	0.95(0.68,1.32)	0.93(0.64,1.34)	351(40.9)	0.72(0.53,0.99)	0.71(0.51,1.00)
Severe illness in the first 2 months	184 (21.4)	1.12(0.76,1.65)	0.90(0.58,1.40)	184(21.4)	1.18(0.82,1.69)	1.03(0.69,1.53)
Birth weight: Normal (>=2500g)	530 (61.3)	1	1	527(61.2)	1	1
Low (<2500g)	45 (5.2)	4.09(2.19,7.61)	4.12(2.06,8.21)	44(5.1)	1.57(0.80,3.05)	1.69(0.81,3.50)
Not measured	289 (33.5)	0.96(0.67,1.37)	0.98(0.66,1.46)	290(33.7)	1.63(1.18,2.24)	1.50(1.06,2.13)

Table 4.27: Predictors of infant undernutrition at 12 months of age in Butajira Birth cohort, Ethiopia

Characteristics considered	*Underweight*			*Stunting*		
	n (%) or Mean(SD)	Crude OR (95% CI)	Adjusted OR (95% CI)	N (%) or Mean(SD)	Crude OR (95% CI)	Adjusted OR (95% CI)
Feeding practices individually adjusted for full model						
Non-exclusive breast-feeding at 2 months	146(15.9)	0.82(0.52,1.29)	0.87(0.53,1.42)	145(15.9)	1.31(0.92,1.87)	1.46(0.97,2.20)
No pre-lacteal food	864(95.4)	1.15(0.52,2.53)	1.02(0.42,2.48)	857(95.2)	0.98(0.53,1.81)	0.99(0.50,1.95)
Colostrums not given	166(18.3)	1.22(0.82,1.82)	1.04(0.66,1.63)	165(18.3)	0.84(0.60,1.18)	0.78(0.53,1.15)
Breast feeding delayed for 1 hour	618(68.1)	1.21(0.85,1.72)	1.20(0.81,1.78)	614(68.2)	0.94(0.71,1.24)	0.85(0.62,1.17)
Fully adjusted model						
Characteristics of mother						
Age (years)	26.9(6.2)	1.04(1.01,1.06)	1.04(1.00,1.08)	26.9(6.2)	1.01(0.99,1.04)	1.02(0.99,1.06)
Height (metres)	1.58(0.06)	1.00(0.97,1.02)	1.00(0.97,1.03)	1.58(0.06)	0.99(0.97,1.01)	0.99(0.97,1.02)
Mid upper arm circumference (cm)	24.7(2.1)	0.91(0.84,0.98)	0.89(0.81,0.97)	24.7(2.1)	0.95(0.89,1.01)	0.95(0.89,1.02)
Being in polygamous marriage	162(17.6)	0.90(0.59,1.38)	0.77(0.45,1.33)	162(17.7)	1.06(0.76,1.49)	1.19(0.76,1.87)
Autonomy scale (0-5)	1.1(1.8)	1.02(0.94,1.11)	1.04(0.95,1.15)	1.1(1.8)	0.97(0.91,1.04)	1.00(0.92,1.08)
Use khat and/or alcohol	104(11.3)	0.70(0.41,1.22)	0.81(0.45,1.47)	104(11.4)	0.77(0.51,1.16)	0.81(0.51,1.28)
Had at least one obstetric complication	573(64.5)	0.91(0.65,1.27)	1.03(0.71,1.49)	571(64.7)	0.86(0.65,1.14)	0.92(0.68,1.26)
Household characteristics						
Urban residence	124(13.5)	0.43(0.24,0.78)	0.38(0.16,0.90)	123(13.4)	0.53(0.36,0.79)	0.52(0.28,0.99)
Number of under 5 children: 0	189(20.5)	1	1	187(20.4)	1	1
1	470(51.0)	1.69(1.07,2.67)	1.33(0.77,2.30)	466(50.9)	1.20(0.85,1.68)	0.96(0.63,1.46)
≥ 2	263(28.5)	1.70(1.04,2.79)	1.75(0.98,3.14)	263(28.7)	1.15(0.79,1.67)	1.19(0.76,1.87)
Age of father in years	36.0(8.8)	1.02(1.00,1.04)	1.02(0.98,1.05)	36.1(8.9)	1.01(0.99,1.02)	1.00(0.97,1.02)
Poverty index (0 to 11)	7.3(2.0)	1.08(0.99,1.17)	0.85(0.74,0.97)	7.3(2.0)	1.08(1.01,1.15)	0.98(0.88,1.09)
Poor sanitary condition (0 to 3)	1.5(1.0)	1.45(1.22,1.72)	1.39(1.13,1.73)	1.5(0.9)	1.37(1.19,1.58)	1.30(1.08,1.55)
Higher level of social support (0 to 4)	1.1(1.1)	1.09(0.94,1.26)	1.07(0.91,1.26	1.1(1.1)	1.08(0.95,1.22)	1.06(0.93,1.22)
Characteristics of index child						
Female gender	456(49.5)	0.80(0.58,1.10)	0.86(0.60,1.23)	449(49.0)	0.48(0.37,0.63)	0.46(0.34,0.62)
Not immunised at two months	381(41.5)	1.22(0.89,1.69)	1.15(0.81,1.65)	379(41.6)	0.96(0.74,1.25)	0.97(0.72,1.30)
Severe illness in the first 2 months	193(21.0)	1.13(0.77,1.66)	1.17(0.76,1.79)	192(21.1)	0.91(0.66,1.26)	0.83(0.58,1.19)
Birth weight: Normal (>=2500gm)	570(61.8)	1	1	568(62.0)	1	1
Low (<2500gm)	43(4.7)	3.08(1.62,5.85)	3.00(1.47,6.13)	41(4.5)	2.71(1.39,5.28)	2.88(1.38,6.03)
Not measured	309(33.5)	1.23(0.88,1.73)	1.28(0.87,1.87)	307(35.5)	1.94(1.46,2.57)	2.18(1.59,3.01)

4.4 Unadjusted and adjusted effects of maternal CMD on infant growth – results from linear regressions and logistic regressions

Length

Unadjusted, partially adjusted and fully adjusted effects of antenatal and postnatal CMD on infant length which is summarized from linear regression analysis are presented in tables 4.28. There was no significant effect of either antenatal or two month postnatal CMD on infant length at two, six or twelve months of age before or after adjusting for potential confounding variables. Regression coefficients and their 95% confidence intervals from bivariate and multivariable regression evaluating the effect of the course of CMD on infant length at two, six and twelve months of age are summarized in table 4.29. There was no significant difference in six or twelve month mean length of infants if their mothers had antenatal CMD but remitted after birth or incident postnatal CMD or chronic CMD (continuously from pregnancy to postnatal period) compared to infants whose mother did not have CMD either in pregnancy or at two month postnatal. However, two month old infants whose mothers had antenatal CMD but remitted after birth were 1.00 cm longer (95%CI: 0.01, 1.98) before adjusting for other risk factors which then slightly increased ($\hat{\beta}$=1.10; 95%CI: 0.02, 2.18) in a fully adjusted model. The latter association was not confounded by infant characteristics ($\hat{\beta}$=1.12; 95%CI: 0.15, 2.09) but it was not significant after adjusting for household characteristics or maternal characteristics or early infant feeding practices.

Table 4.28: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on length score of infant at selected follow-up time points

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Antenatal	0.62(-0.26, 1.49)	0.19(-0.53, 0.91)	0.01(-0.74, 0.77)
Postnatal	-1.04(-2.39, 0.30)	0.56(-0.54, 1.65)	0.13(-1.05, 1.31)
Adjusted for household characteristics			
Antenatal	0.61(-0.30, 1.52)	0.18(-0.56, 0.93)	0.41(-0.36, 1.18)
Postnatal	-0.95(-2.31, 0.41)	0.52(-0.59, 1.63)	0.11(-1.07, 1.29)
Maternal characteristics			
Antenatal	0.60(-0.32, 1.52)	0.22(-0.54, 0.97)	0.15(-0.64, 0.94)
Postnatal	-0.89(-2.25, 0.46)	0.79(-0.32, 1.90)	0.24(-0.95, 1.43)
Infant Characteristics			
Antenatal	0.69(-0.18, 1.56)	0.11(-0.60, 0.82)	-0.09(-0.84, 0.66)
Postnatal	-1.17(-2.51, 0.18)	0.34(-0.76, 1.43)	-0.20(-1.38, 0.97)
Feeding practices			
Antenatal	0.52(-0.37, 1.41)	0.11(-0.63, 0.85)	-0.07(-0.84, 0.71)
Postnatal	-1.06(-2.40, 0.27)	0.59(-0.51, 1.69)	0.13(-1.04, 1.31)
Fully adjusted			
Antenatal	0.67(-0.28, 1.63)	0.23(-0.56, 1.02)	0.45(-0.37, 1.26)
Postnatal	-0.86(-2.23, 0.51)	0.65(-0.47, 1.78)	-0.07(-1.26, 1.12))

Table 4.29: Unadjusted, partially adjusted and fully adjusted effect of course of CMD on length of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Never exposed			
Pregnancy only	1.00(0.01, 1.98)	0.05(-0.76, 0.86)	-0.12(-0.96, 0.72)
Postnatal only	-1.16(-3.19, 0.87)	0.34(-1.38, 2.05)	-0.39(-2.15, 1.37)
Both time points	-0.79(-2.55, 0.98)	0.71(-0.70, 2.11)	0.51(-1.05, 2.08)
Adjusted for household characteristics			
Never exposed			
Pregnancy only	0.99(-0.03, 2.02)	0.01(-0.84, 0.86)	0.33(-0.54, 1.20)
Postnatal only	-1.05(-3.09, 1.00)	0.24(-1.49, 1.96)	-0.51(-2.26, 1.24)
Both time points	-0.72(-2.50, 1.06)	0.71(-0.71, 2.13)	0.65(-0.90, 2.20)
Maternal characteristics			
Never exposed			
Pregnancy only	0.96(-0.08, 1.99)	0.00(-0.85, 0.86)	-0.01(-0.89, 0.88)
Postnatal only	-0.93(-2.96, 1.09)	0.60(-1.12, 2.33)	-0.29(-2.06, 1.48)
Both time points	-0.67(-2.45, 1.11)	0.91(-0.51, 2.34)	0.66(-0.92, 2.24)
Infant Characteristics			
Never exposed			
Pregnancy only	1.12(0.15, 2.09)	0.05(-0.75, 0.85)	-0.11(-0.95, 0.72)
Postnatal only	-1.19(-3.20, 0.82)	0.34(-1.35, 2.03)	-0.43(-2.17, 1.32)
Both time points	-0.96(-2.73, 0.81)	0.34(-1.07, 1.75)	-0.05(-1.61, 1.52)
Feeding practices			
Never exposed			
Pregnancy only	0.88(-0.13, 1.88)	-0.09(-0.93, 0.76)	-0.24(-1.11, 0.63)
Postnatal only	-1.29(-3.32, 0.73)	0.32(-1.41, 2.04)	-0.37(-2.13, 1.39)
Both time points	-0.75(-2.50, 1.00)	0.76(-0.65, 2.17)	0.49(-1.07, 2.05)
Fully adjusted			
Never exposed			
Pregnancy only	1.10(0.02, 2.18)	0.06(-0.85, 0.97)	0.50(-0.47, 1.42)
Postnatal only	-0.81(-2.83, 1.22)	0.55(-1.18, 2.28)	-0.32(-2.07, 1.42)
Both time points	-0.69(-2.48, 1.10)	0.73(-0.72, 2.17)	0.24(-1.34, 1.81)

Weight

Regression coefficients and their 95% confidence intervals from bivariate and multivariable regression models evaluating the effect antenatal and postnatal CMD on infant weight measured at two, six or twelve months of age are summarized in Table 3.30. In bivariate analysis postnatal CMD was positively associated with infant weight at six months ($\hat{\beta}$ =0.39kg; 95%CI: 0.05 to 0.72) and in a fully adjusted model the effect size were relatively reduced becoming marginally non-significant ($\hat{\beta}$ =0.31kg; 95%CI: -0.01 to 0.63). Partially adjusting the association did not show significant confounding effects of household characteristics, maternal characteristics or early infant feeding practices of the mother.

Table 4.30: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on weight of infant at selected follow-up time points

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Antenatal	0.12(-0.05, 0.30)	-0.03(-0.24, 0.19)	-0.05(-0.30, 0.20)
Postnatal	0.26(-0.01, 0.53)	0.39(0.05, 0.72)	0.22(-0.17, 0.61)
Adjusted for			
Household characteristics			
Antenatal	0.13(-0.05, 0.31)	0.03(-0.19, 0.24)	0.06(-0.20, 0.31)
Postnatal	0.24(-0.03, 0.52)	0.34(0.02, 0.67)	0.21(-0.17, 0.59)
Maternal characteristics			
Antenatal	0.12(-0.06, 0.30)	-0.02(-0.25, 0.20)	-0.00(-0.26, 0.26)
Postnatal	0.28(0.01, 0.55)	0.42(0.08, 0.75)	0.25(-0.13, 0.63)
Infant Characteristics			
Antenatal	0.14(-0.04, 0.31)	-0.04(-0.25, 0.17)	-0.06(-0.31, 0.19)
Postnatal	0.20(-0.07, 0.46)	0.32(-0.01, 0.65)	0.13(-0.26, 0.52)
Feeding practices			
Antenatal	0.09(-0.09, 0.28)	-0.05(-0.27, 0.17)	-0.07(-0.33, 0.18)
Postnatal	0.25(-0.02, 0.52)	0.37(0.03, 0.70)	0.21(-0.18, 0.59)
Fully adjusted			
Antenatal	0.11(-0.08, 0.29)	0.03(-0.19, 0.26)	0.11(-0.15, 0.37)
Postnatal	0.22(-0.05, 0.48)	0.31(-0.01, 0.63)	0.16(-0.21, 0.54)

Table 4.31: Unadjusted, partially adjusted and fully adjusted effect of course of CMD on weight of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Never exposed			
Pregnancy only	0.08(-0.11, 0.28)	-0.15(-0.39, 0.09)	-0.21(-0.49, 0.07)
Postnatal only	0.22(-0.19, 0.62)	0.33(-0.19, 0.86)	-0.23(-0.81, 0.36)
Both time points	0.31(-0.05, 0.67)	0.40(-0.03, 0.82)	0.52(0.01, 1.03)
Adjusted for			
Household characteristics			
Never exposed			
Pregnancy only	0.08(-0.12, 0.28)	-0.10(-0.34, 0.15)	-0.11(-0.39, 0.17)
Postnatal only	0.18(-0.22, 0.59)	0.21(-0.29, 0.72)	-0.29(-0.86, 0.28)
Both time points	0.31(-0.05, 0.66)	0.42(0.00, 0.83)	0.57(0.07, 1.06)
Maternal characteristics			
Never exposed			
Pregnancy only	0.07(-0.13, 0.28)	-0.15(-0.40, 0.10)	-0.17(-0.46, 0.12)
Postnatal only	0.28(-0.12, 0.67)	0.39(-0.13, 0.91)	-0.18(-0.75, 0.39)
Both time points	0.29(-0.06, 0.65)	0.41(-0.02, 0.84)	0.54(0.04, 1.05)
Infant Characteristics			
Never exposed			
Pregnancy only	0.12(-0.07, 0.31)	-0.13(-0.37, 0.11)	-0.19(-0.46, 0.09)
Postnatal only	0.20(-0.19, 0.60)	0.31(-0.20, 0.82)	-0.24(-0.82, 0.33)
Both time points	0.21(-0.14, 0.56)	0.30(-0.12, 0.72)	0.38(-0.13, 0.89)
Feeding practices			
Never exposed			
Pregnancy only	0.03(-0.17, 0.24)	-0.20(-0.45, 0.05)	-0.27(-0.56, 0.02)
Postnatal only	0.16(-0.24, 0.57)	0.24(-0.28, 0.77)	-0.27(-0.87, 0.29)
Both time points	0.33(-0.03, 0.69)	0.42(-0.01, 0.84)	0.53(0.03, 1.04)
Fully adjusted			
Never exposed			
Pregnancy only	0.07(-0.14, 0.28)	-0.08(-0.33, 0.18)	-0.03(-0.32, 0.27)
Postnatal only	0.20(-0.19, 0.59)	0.20(-0.29, 0.70)	-0.27(-0.82, 0.29)
Both time points	0.24(-0.11, 0.59)	0.37(-0.04, 0.78)	0.49(-0.00, 0.98)

Results from bivarite and multivariable regression analysis that takes course of CMD as main exposure variable and weigh of infants at two, six and twelve months as outcome variables are summarized in table 4.31. There was no significant association between the course of CMD and weight of infants measured at two or six months of age. At twelve months of age infants whose mothers had chronic CMD were 0.52kg ($\hat{\beta}$=0.52kg; 95%CI: 0.01 to 1.03) heavier than infants whose mother did not have CMD either during

third trimester of pregnancy or during two month postnatal follow-up. After adjusting for potential confounding variables the change in the effect size was minimal but it was marginally non-significant ($\hat{\beta}$=0.49kg; 95%CI: -0.00 to 0.98)

Length-for-age

In bivariate and multivariate regression analysis antenatal or postnatal CMD were not significantly associated with six or twelve month old infants' length-for-age z (Table 4.32.). At two months of age postnatal CMD had significant negative effect on length-for-age z ($\hat{\beta}$ = - 0.62 ; 95%CI: -1.14 to -0.10) which turned out to be non-significant in a fully adjusted model ($\hat{\beta}$ = - 0.42 ; 95%CI: -0.95 to 0.11). Partially adjusting for household characteristics, maternal characteristics, infant characteristics or early infant feeding practices of the mother did not affect statistical significance of the negative effect of postnatal CMD.

Table 4.32: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on length-for-age z of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Antenatal	0.17(-0.17, 0.59)	-0.10(-0.43, 0.23)	-0.05(-0.36, 0.25)
Postnatal	-0.62(-1.14, -0.10)	-0.07(-0.57, 0.44)	-0.12(-0.60, 0.36)
Adjusted for			
Household characteristics			
Antenatal	0.21(-0.14, 0.56)	-0.10(-0.45, 0.24)	0.11(-0.20, 0.43)
Postnatal	-0.55(-1.07, -0.03)	-0.08(-0.59, 0.44)	-0.14(-0.61, 0.34)
Maternal characteristics			
Antenatal	0.13(-0.22, 0.49)	-0.08(-0.43, 0.27)	-0.01(-0.33, 0.31)
Postnatal	-0.59(-1.12, -0.07)	0.06(-0.46, 0.57)	-0.06(-0.54, 0.42)
Infant Characteristics			
Antenatal	0.21(-0.13, 0.55)	-0.08(-0.40, 0.34)	-0.04(-0.34, 0.36)
Postnatal	-0.55(-1.08, -0.03)	0.01(-0.49, 0.51)	-0.06(-0.53, 0.42)
Feeding practices			
Antenatal	0.19(-0.16, 0.54)	-0.14(-0.48, 0.20)	-0.08(-0.39, 0.24)
Postnatal	-0.61(-1.13, -0.08)	-0.05(-0.56, 0.46)	-0.12(-0.60, 0.36)
Fully adjusted			
Antenatal	0.28(-0.09, 0.64)	-0.03(-0.39, 0.33)	0.19(-0.13, 0.52)
Postnatal	-0.42(-0.95, 0.11)	0.18(-0.33, 0.70)	-0.01(-0.48, 0.46)

Table 4.33: Unadjusted, partially adjusted and fully adjusted effect of course of CMD on length-for-age z at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Never exposed			
Pregnancy only	0.29(-0.09, 0.66)	-0.09(-0.47, 0.28)	-0.06(-0.49, 0.28)
Postnatal only	-0.89(-1.66, -0.13)	-0.04(-0.84, 0.77)	-0.22(-0.94, 0.49)
Both time points	-0.35(-1.04, 0.35)	-0.10(-0.74, 0.54)	-0.04(-0.68, 0.59)
Adjusted for			
Household characteristics			
Never exposed			
Pregnancy only	0.32(-0.07, 0.71)	-0.11(-0.50, 0.28)	0.14(-0.21, 0.48)
Postnatal only	-0.84(-1.61, -0.08)	-0.07(-0.88, 0.75)	-0.30(-1.00, 0.41)
Both time points	-0.26(-0.95, 0.44)	-0.10(-0.75, 0.55)	0.01(-0.61, 0.64)
Maternal characteristics			
Never exposed			
Pregnancy only	0.23(-0.17, 0.62)	-0.11(-0.51, 0.28)	-0.02(-0.38, 0.33)
Postnatal only	-0.88(-1.65, -0.12)	0.08(-0.73, 0.90)	-0.18(-0.89, 0.54)
Both time points	-0.30(-1.00, 0.39)	0.02(-0.63, 0.68)	0.03(-0.61, 0.67)
Infant Characteristics			
Never exposed			
Pregnancy only	0.30(-0.07, 0.68)	-0.09(-0.45, 0.27)	-0.07(-0.40, 0.27)
Postnatal only	-0.87(-1.63, -0.10)	0.07(-0.72, 0.86)	-0.19(-0.88, 0.51)
Both time points	-0.23(-0.94, 0.47)	-0.04(-0.68, 0.60)	0.04(-0.59, 0.66)
Feeding practices			
Never exposed			
Pregnancy only	0.32(-0.07, 0.71)	-0.15(-0.54, 0.24)	-0.09(-0.45, 0.26)
Postnatal only	-0.88(-1.66, -0.11)	-0.04(-0.85, 0.77)	-0.22(-0.94, 0.49)
Both time points	-0.32(-1.00, 0.37)	-0.08(-0.73, 0.56)	-0.05(-0.69, 0.58)
Fully adjusted			
Never exposed			
Pregnancy only	0.37(-0.05, 0.78)	-0.09(-0.50, 0.33)	0.20(-0.16, 0.57)
Postnatal only	-0.72(-1.48, 0.04)	0.21(-0.60, 1.02)	-0.16(-0.85, 0.54)
Both time points	-0.09(-0.80, 0.61)	0.15(-0.50, 0.81)	0.14(-0.48, 0.77)

Course of CMD was not significantly associated with infant length-for-age at six or twelve months of age (Table 4.33). At two months of age incident postnatal CMD had significant negative effect on length-for-age Z ($\hat{\beta}$ = - 0.89 ; 95%CI: -1.66 to -0.13) in the unadjusted model but not in a fully adjusted model ($\hat{\beta}$ = - 0.72 ; 95%CI: -1.48 to 0.04). The negative effect of incident postnatal CMD on length-for-age z at two months follow-

up was not confounded by household characteristics, maternal characteristics, infant characteristics or early infant feeding practices of the mother.

Weight-for-age

Regression coefficients and corresponding 95% confidence intervals from bivariate and multivariable regression analysis assessing the effects of maternal CMD on infants' weight-for-age z score at 2, 6 and 12 months of age are summarized in Tables 4.34 and 4.35. The only significant effects were observed in fully adjusted models where (a) prevalent postnatal CMD had positive effect on two month old infant's weight-for-age z ($\hat{\beta}$ = 0.40 ; 95%CI: 0.05 to 0.76) and (b) chronic CMD compared to no postnatal and antenatal CMD had positive effect on weight-for-age z at the age of two months ($\hat{\beta}$ = 0.52 ; 95%CI: 0.05 to 0.99) and at the age of twelve months ($\hat{\beta}$ = 0.54 ; 95%CI: 0.03 to 1.05) .

Table 4.34: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on weight-for-age z at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Antenatal	0.15(-0.08, 0.39)	-0.12(-0.39, 0.15)	-0.12(-0.38, 0.15)
Postnatal	0.32(-0.04, 0.68)	0.26(-0.15, 0.68)	0.14(-0.26, 0.54)
Adjusted for Household characteristics			
Antenatal	0.17(-0.06, 0.41)	-0.05(-0.31, 0.22)	-0.01(-0.27, 0.25)
Postnatal	0.29(-0.07, 0.64)	0.20(-0.20, 0.60)	0.11(-0.28, 0.51)
Maternal characteristics			
Antenatal	0.13(-0.11, 0.37)	-0.11(-0.39, 0.17)	-0.07(-0.33, 0.20)
Postnatal	0.35(-0.01, 0.71)	0.31(-0.10, 0.73)	0.18(-0.22, 0.57)
Infant Characteristics			
Antenatal	0.19(-0.04, 0.41)	-0.08(-0.34,0.18)	-0.10(-0.35, 0.16)
Postnatal	0.35(0.01, 0.70)	0.36(-0.04, 0.77)	0.18(-0.22, 0.57)
Feeding practices			
Antenatal	0.12(-0.11, 0.36)	-0.16(-0.43, 0.12)	-0.15(-0.41, 0.12)
Postnatal	0.31(-0.04, 0.67)	0.24(-0.17, 0.65)	-0.12(-0.28, 0.52)
Fully adjusted			
Antenatal	0.17(-0.07, 0.42)	0.02(-0.26, 0.30)	0.09(-0.18, 0.36)
Postnatal	0.40(0.05, 0.76)	0.36(-0.05, 0.76)	0.20(-0.18, 0.59)

Table 4.35: Unadjusted, partially adjusted and fully adjusted effect of course of CMD on weight-for-age z of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months* $\hat{\beta}$ *(95% CI)*	*6 months* $\hat{\beta}$ *(95%CI)*	*12 months* $\hat{\beta}$ *(95%CI)*
Unadjusted			
Never exposed			
Pregnancy only	0.09(-0.17, 0.35)	-0.19(-0.49, 0.11)	-0.26(-0.55, 0.02)
Postnatal only	0.21(-0.32, 0.75)	0.39(-0.26, 1.03)	-0.23(-0.84, 0.37)
Both time points	0.42(-0.05, 0.89)	0.15(-0.37, 0.68)	0.37(-0.15, 0.90)
Adjusted for household characteristics			
Never exposed			
Pregnancy only	0.11(-0.16, 0.37)	-0.11(-0.42, 0.19)	-0.15(-0.44, 0.15)
Postnatal only	0.14(-0.39, 0.67)	0.24(-0.39, 0.87)	-0.32(-0.92, 0.27)
Both time points	0.42(-0.05, 0.89)	0.17(-0.35, 0.68)	0.42(-0.09, 0.93)
Maternal characteristics			
Never exposed			
Pregnancy only	0.06(-0.22, 0.33)	-0.19(-0.50, 0.12)	-0.22(-0.51, 0.08)
Postnatal only	0.28(-0.25, 0.81)	0.46(-0.18, 1.11)	-0.19(-0.78, 0.41)
Both time points	0.42(-0.06, 0.89)	0.18(-0.34, 0.71)	0.41(-0.11, 0.93)
Infant Characteristics			
Never exposed			
Pregnancy only	0.12(-0.13, 0.37)	-0.17(-0.46, 0.13)	-0.25(-0.53, 0.04)
Postnatal only	0.24(-0.27, 0.76)	0.48(-0.15, 1.11)	-0.21(-0.80, 0.39)
Both time points	0.46(0.00, 0.93)	0.26(-0.26, 0.78)	0.43(-0.10, 0.95)
Feeding practices			
Never exposed			
Pregnancy only	0.04(-0.23, 0.31)	-0.26(-0.57, 0.05)	-0.32(-0.62, -0.02)
Postnatal only	0.16(-0.38, 0.69)	0.28(-0.37, 0.92)	-0.29(-0.89, 0.31)
Both time points	0.44(-0.03, 0.92)	0.18(-0.34, 0.71)	0.38(-0.14, 0.91)
Fully adjusted			
Never exposed			
Pregnancy only	0.07(-0.21, 0.35)	-0.08(-0.40, 0.24)	-0.06(-0.82, 0.33)
Postnatal only	0.27(-0.25, 0.79)	0.37(-0.24, 0.99)	-0.24(-0.82, 0.33)
Both time points	0.52(0.05, 0.99)	0.33(-0.18, 0.84)	0.54(0.03, 1.05)

Stunting and underweight

The results of bivariate and multivariable logistic regression with antenatal CMD (prevalent cases) and postnatal CMD (prevalent cases) as the main exposures for infant stunting and underweight are presented in Tables 4.36 and 4.37. There was no significant effect of either antenatal or postnatal CMD upon prevalence of infant stunting and underweight at either time point, both before and after adjusting for potential confounding variables. In a multivariable logistic regression use of SRQ score as a continuous exposure variable did not altered our finding of no association between CMD and prevalence of infant stunting or underweight.

Table 4.36: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on stunting of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months OR (95% CI)*	*6 months OR (95%CI)*	*12 months OR (95%CI)*
Unadjusted			
Antenatal	0.91(0.50, 1.66)	0.91(0.56, 1.46)	1.13(0.76, 1.70)
Postnatal	1.26(0.54, 2.90)	0.94(0.45, 1.97)	0.84(0.45, 1.58)
Adjusted for household characteristics			
Antenatal	0.72(0.37, 1.42)	0.90(0.55, 1.47)	1.00(0.65, 1.52)
Postnatal	1.09(0.46, 2.60)	0.93(0.44, 1.95)	0.88(0.46, 1.68)
Maternal characteristics			
Antenatal	0.94(0.50, 1.76)	0.84(0.50, 1.39)	1.10(0.72, 1.69)
Postnatal	1.28(0.55, 3.00)	0.85(0.40, 1.80)	0.80(0.42, 1.52)
Infant Characteristics			
Antenatal	0.86(0.47, 1.58)	0.96(0.59, 1.56)	1.18(0.78, 1.79)
Postnatal	1.10(0.47, 2.58)	0.91(0.44, 1.91)	0.82(0.44, 1.55)
Feeding practices			
Antenatal	0.92(0.50, 1.68)	0.96(0.59, 1.57)	1.18(0.78, 1.78)
Postnatal	1.25(0.54, 2.89)	0.92(0.44, 1.92)	0.81(0.43, 1.53)
Fully adjusted			
Antenatal	0.68(0.33, 1.39)	0.80(0.46, 1.38)	1.00(0.62, 1.60)
Postnatal	1.02(0.42, 2.47)	0.66(0.30, 1.45)	0.80(0.40, 1.59)

OR = Odds ratio

Table 4.37: Unadjusted, partially adjusted and fully adjusted effect of antenatal and postnatal prevalent CMD on underweight of infant at 2, 6 and 12 months of age

Model/timing for main exposure	*2 months OR (95% CI)*	*6 months OR (95%CI)*	*12 months OR (95%CI)*
Unadjusted			
Antenatal	0.65(0.31, 1.39)	1.43(0.89,2.30)	1.34(0.84, 2.12)
Postnatal	0.20(0.03, 1.45)	0.66(0.27, 1.61)	0.61(0.25, 1.47)
Adjusted for			
Household characteristics			
Antenatal	0.64(0.29, 1.43)	1.37(0.83, 2.27)	1.13(0.69, 1.84)
Postnatal	0.22(0.03, 1.61)	0.76(0.31, 1.87)	0.63(0.26, 1.54)
Maternal characteristics			
Antenatal	0.80(0.37, 1.72)	1.50(0.91, 2.48)	1.17(0.71, 1.92)
Postnatal	0.22(0.03, 1.63)	0.60(0.24, 1.47)	0.53(0.22, 1.29)
Infant Characteristics			
Antenatal	0.60(0.28, 1.29)	1.50(0.92, 2.46)	1.33(0.82, 2.14)
Postnatal	0.17(0.02, 1.24)	0.68(0.28, 1.65)	0.62(0.26, 1.49)
Feeding practices			
Antenatal	0.70(0.33, 1.49)	1.52(0.93, 2.49)	1.32(0.82, 2.13)
Postnatal	0.20(0.03, 1.46)	0.67(0.28, 1.65)	0.61(0.25, 1.47)
Fully adjusted			
Antenatal	0.66(0.28, 1.55)	1.28(0.73, 2.24)	0.81(0.46, 1.43)
Postnatal	0.18(0.02, 1.37)	0.56(0.22, 1.46)	0.52(0.21, 1.32)

OR= Odds ratio

Odds ratios and corresponding 95% confidence intervals from bivariate and multivariable logistic regressions assessing the association between the course of CMD from pregnancy to two months postnatally (a four level categorical variable) and prevalence of infant stunting and underweight are presented in tables 4.38 and 4.39. The reference category for this exposure was those mothers who had low levels of CMD at both assessment points. Prior to adjustment for possible confounding factors, infants whose mothers had high levels of CMD during pregnancy which resolved after delivery were more likely to be underweight at 12 months of age (OR = 1.71; 95% CI: 1.05 – 2.80), with a non-significant trend in the same direction at six months (OR=1.53; 95% CI: 0.91-2.60) and for stunting at 12 months (OR = 1.30, 95% CI: 0.83 - 2.03). The excess risk for an infant being underweight at 12 months of age remained significant after adjusting for infant characteristics and early infant feeding practices of the mother but became statistically non-significant after adjusting for maternal characteristics or household characteristics. Although the risk for underweight at six months and for stunting at

twelve months was not statistically significantly associated with antenatal CMD which resolved after delivery , a consistent trend in the same direction still remained after adjusting for each group of confounding variables. In the final multivariable model, adjusting for all of the potential confounders simultaneously, the course of CMD was not significantly associated with infant nutritional status at either two, six or twelve months of age.

Table 4.38: Unadjusted, partially adjusted and fully adjusted effects of course of CMD on stunting of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months OR (95% CI)*	*6 months OR (95%CI)*	*12 months OR (95%CI)*
Unadjusted			
Pregnancy only	1.03(0.54, 1.97)	0.90(0.52, 1.56)	1.30(0.83, 2.03)
Postnatal only	2.28(0.80, 6.53)	0.99(0.31, 3.13)	1.10(0.43, 2.80)
Both time points	0.59(0.14, 2.58)	0.90(0.35, 2.31)	0.71(0.30, 1.65)
Adjusted for			
Household characteristics			
Pregnancy only	0.83(0.40, 1.73)	0.90(0.51, 1.58)	1.12(0.70, 1.79)
Postnatal only	2.03(0.68, 6.06)	0.97(0.30, 3.09)	1.26(0.28, 3.26)
Both time points	0.50(0.11, 2.20)	0.89(0.35, 2.29)	0.68(0.29, 1.62)
Maternal characteristics			
Pregnancy only	1.08(0.55, 2.14)	0.84(0.47, 1.50)	1.27(0.79, 2.04)
Postnatal only	2.35(0.81, 6.83)	0.87(0.27, 2.78)	1.04(0.40, 2.26)
Both time points	0.60(0.14, 2.65)	0.82(0.32, 2.14)	0.69(0.29, 1.63)
Infant Characteristics			
Pregnancy only	1.03(0.54, 1.97)	0.98(0.56, 1.71)	1.39(0.87, 2.21)
Postnatal only	2.23(0.77, 6.44)	0.98(0.31, 3.14)	1.08(0.42, 2.75)
Both time points	0.47(0.11, 2.07))	0.87(0.34, 2.23)	0.70(0.30, 1.64)
Feeding practices			
Pregnancy only	1.06(0.55, 2.03)	0.98(0.56, 1.71)	1.39(0.87, 2.21)
Postnatal only	2.40(0.83, 6.92)	0.98(0.31, 3.14)	1.08(0.42, 2.75)
Both time points	0.57(0.13, 2.48)	0.87(0.34, 2.23)	0.70(0.30, 1.64)
Fully adjusted			
Pregnancy only	0.84(0.38, 1.86)	0.86(0.46, 1.62)	1.14(0.67, 1.96)
Postnatal only	2.11(0.69, 6.49)	0.67(0.20, 2.27)	1.06(0.39, 2.94)
Both time points	0.41(0.09, 1.87)	0.64(0.24, 1.73)	0.66(0.27, 1.63)

OR = Odds ratio

Table 4.39: Unadjusted, partially adjusted and fully adjusted effect of course of CMD on underweight of infant at two, six and 12 months of age

Model/timing for main exposure	*2 months OR (95% CI)*	*6 months OR (95%CI)*	*12 months OR (95%CI)*
Unadjusted			
Pregnancy only	0.72(0.32, 1.60)	1.53(0.91, 2.60)	1.71(1.05, 2.80)
Postnatal only	-	0.26(0.03, 2.02)	1.11(0.36, 3.42)
Both time points	0.35(0.05, 2.62)	1.03(0.38, 2.81)	0.35(0.08, 1.52)
Adjusted for household characteristics			
Pregnancy only	0.70(0.29, 1.66)	1.44(0.82, 2.54)	1.46(0.86, 2.47)
Postnatal only	-	0.33(0.04, 2.57)	1.31(0.41, 4.17)
Both time points	0.38(0.05, 2.87)	1.09(0.40, 3.03)	0.32(0.07, 1.40)
Maternal characteristics			
Pregnancy only	0.88(0.39, 2.00)	1.65(0.95, 2.87)	1.54(0.91, 2.63)
Postnatal only	-	0.23(0.03, 1.78)	0.99(0.32, 3.13)
Both time points	0.41(0.05, 3.17)	0.98(0.35, 2.73)	0.30(0.07, 1.30)
Infant Characteristics			
Pregnancy only	0.67(0.29, 1.54)	1.68(0.97, 2.91)	1.75(1.05, 2.94)
Postnatal only	-	0.30(0.04, 2.34)	1.17(0.38, 3.61)
Both time points	0.29(0.04, 2.25)	0.98(0.36, 2.70)	0.35(0.08, 1.49)
Feeding practices			
Pregnancy only	0.78(0.35, 1.76)	1.68 (0.97, 2.91)	1.75(1.05, 2.93)
Postnatal only	-	0.30(0.04, 2.34)	1.17(0.38, 3.61)
Both time points	0.35(0.05, 2.60)	0.98(0.36, 2.70)	0.35(0.08, 149)
Fully adjusted			
Pregnancy only	0.75(0.30, 1.90)	1.43(0.76, 2.71)	1.07(0.58, 1.97)
Postnatal only	-	0.21(0.02, 1.86)	1.07(0.33, 3.62)
Both time points	0.34(0.04, 2.64)	0.85(0.29, 2.50)	0.25(0.06, 1.15)

OR = Odds ratio

4.5 Summary of the chapter

CMD and infant growth

Before adjusting for other risk factors but not in the multivariable analysis prevalent postnatal and incident postnatal CMD symptoms compared to absence of perinatal CMD symptoms had a negative effect on length-for-age z at two month. These associations were also implicated in a non-significant trend of the effect of prevalent and incident postnatal CMD as being risk factors of stunting and for having shorter length at two months of age. At six months of age there was no significant association of CMD and infant growth except positive effect of postnatal CMD (i.e. prevalent) on infant weight which was not significant in multivariable analysis. Postnatal CMD (i.e. prevalent) was also significantly associated with an increase in weight-for-age z at two months after adjusting for other risk factors which was not otherwise significant in bivariate analysis.

CMD during pregnancy which was resolved after birth was significantly associated with the risk of underweight at twelve months in bivariate analysis and with better performance in length at two months before and after adjusting for other risk factors. Crude and adjusted result also showed non-significant trend in risk of underweight at six months associated with prevalent antenatal CMD and antenatal CMD which was resolved after birth. In bivariate analysis persistent CMD was associated with having increased mean weight at twelve months and in multivariable analysis it was associated with increased weight-for-age z at two and twelve months of age.

Other risk factors of infant growth

Reduced maternal MUAC, rural residence, poor sanitary condition of the household, male gender and low birth weight were significant risk factors for compromised infant growth at 2, 6 and 12 months of age. At twelve months of age these associations were significant across all the six growth outcomes except that gender and MUAC were not associated with underweight and stunting, respectively. When infants were six months old (a) only male gender was significant risk factor associated with stunting, (b) poor sanitary condition had only significant effect on weight, weight-for-age and underweight and, (c) rural residence was not significantly associated with lighter weight of infants. At two months birth weight was not significantly associated with length-for-age and stunting, poor sanitary condition of the household was not significantly associated with

weight and underweight, rural residence was not a significant predictor of lighter weight, low weight-for-age z and underweight, and MUAC was only significantly associated with weight and weight-for-age z.

The negative effect of increased parental age was not observed in two months, it was marginally significant or significant in bivaraite analysis at six months with three weight derived outcomes, and was significant in bivariate or multivariable analysis at twelve months except with length-for-age. In bivariate analysis the poverty index was negatively associated with weight, weight-for-age and underweight at six and twelve months of age. At two months of age scoring high on poor sanitary condition was associated with short length before adjusting for other risk factors but it was not significant in multivariable analysis. At six months, however, there was no association in bivariate analysis and positive association after adjusting for other factors. Infants whose mothers had at least one obstetric complication during delivery had a tendency of performing well (a) in their weight, weight-for-age and of not being underweight at two months, (b) on their weight at six months and (c) on their weight and weight-for-age at twelve months.

Early infant feeding practices of the mothers were not significantly associated with growth outcomes to the extent that they could substantially influence the overall growth of infants. In bivariaite analysis delayed initiation of breast feeding had a tendency of affecting infant weight at two and six months of age and denial of colostrums was associated positively with length-for-age at two months and negatively associated with weight and weight-for-age at 6 and twelve months of age, respectively. In multivariable analysis the only significant associations were the negative effect of delayed initiations of breast feeding on weight-for-age at six months and positive effect of denial of colostrums on length-for-age at two months of age.

CHAPTER 5: RESULTS - UNCONDITIONAL LGMS

5.1 Summary of model description and modelling strategies

The three unconditional LGM (Bollen and Curran 2006) described in chapter three of this thesis and presented below taking length as one example of an outcome variable were fitted to the repeated measurement of growth of infants from 2 to 18 months of age the rural villages of Butajira. The modelling started with the simplest model and moved step by step to the more complex model. Each model has two groups of equations which defines growth of individual over time (level 1) and growth at a group level or average growth (level 2).

Model 1: (Linear)

$$length_{it} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it} \quad \ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots \quad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \quad \beta_i = \mu_\beta + \zeta_{\beta_i} \quad \ldots\ldots\ldots \quad \text{Level 2}$$

Model 2: (Quadratic)

$$length_{it} = \alpha_i + \lambda_t \beta_{1i} + \lambda_t^2 \beta_{2i} + \varepsilon_{it} \quad \ldots\ldots\ldots\ldots \quad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \quad \beta_{1i} = \mu_{\beta_1} + \zeta_{\beta_{1i}} \quad and \quad \beta_{2i} = \mu_{2\beta_2} + \zeta_{\beta_{2i}} \ldots \quad \text{Level 2}$$

Model 3: (Non-linear)

$$length_{it} = \alpha_i + \lambda_t \beta_i + \varepsilon_{it} \quad \ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots \quad \text{Level 1}$$

$$\alpha_i = \mu_\alpha + \zeta_{\alpha_i} \quad and \quad \beta_i = \mu_\beta + \zeta_{\beta_i} \quad \ldots\ldots\ldots \quad \text{Level 2}$$

In these models the values of λ_t determines how time of measurement or passage of time should be coded. In linear and quadratic latent growth models (i.e. models 1 and 2) the values of λ_t have values equal to (t-2) where t corresponds to the time of each measurement and in model 3 two of the λ_t's are fixed to 0 and 1, respectively, and all the rest are estimated freely from the data. This way of coding time guarantees that the growth within an individual is linear (i.e. model 1), quadratic (model 2) or non-linear (model 3) over time. Subtraction of two months from the time of assessment assigns two month assessment to the starting value for growth and gives meaning to the intercept term (i.e. the expected value of α_i) in the regression equations. Although any two time points can be used as reference interval for the change in model 3 we took two month time point as starting time point so that $\lambda_1 = 0$, eighteen month assessment as the point of

final assessment ($\lambda_5 = 1$) and we left all the remaining measurement occasions (i.e. $\lambda_2, \lambda_3, \lambda_4$) to be estimated as free model parameters. The way we parametrize model 3 enable us to interpret any change between the starting point (i.e. two month time point in our case) and the time under investigation as the proportion of change over the whole follow-up time .

In all the above level 1 models we considered two scenarios: (a) all residuals assumed to be mutually independent to enable as to select appropriate growth model for the description of the data and (b) serial correlation allowed between measurement errors of consecutive time points to further improve the fit of the selected level 1 model. In the latter case a given serial correlation was retained in a model if it was statistically significant and did not influence the validity of model parameters.

5.2 Best fitting unconditional LGM for growth outcomes

5.2.1 Length of infants

Based on four goodness of fit indices presented in table 5.1 **a** non-linear LGM that allows serial correlation between residuals at six, nine and twelve months of age gives acceptable description of infant length ($\chi^2(df = 10) = 18.7$, p=0.044; TLI = 0.980, CFI = 0.991, RMSEA = 0.029 with 90%CI: 0.005 to 0.049). In this non-linear LGM time is coded in such a way that gain of length between two months of age and any subsequent age (eg. six months) be scaled relative to the time span between two and eighteen months of age (i.e. sixteen months).

Table 5.1 Fit indices for linear, quadratic and non-linear Latent Growth Models in modelling length of infants between two and 18 months of age in Butajira

Model description	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA(90%CI)***
Model 1: Independent residuals	497.1(16; <0.000)	0.481	0.319	0.170(0.157,0.183)
Model 1: Serial correlation of residuals allowed (**e2&e4; e6&e9; e9&e12**)	417.3(13; <0.000)	0.564	0.296	0.173(0.158,0.187)
Model 2: Independence of residuals	118.0(12; <0.001)	0.886	0.800	0.092(0.077,0.107)
Model 2: Serial correlation of residuals allowed (**e6&e9; e9&e12**)	100.3(10; <0.000)	0.903	0.795	0.093(0.077,0.110)
Model 3: Independence of residuals allowed	47.1(12; <0.000)	0.962	0.934	0.053(0.038,0.069)
Model 3: Serial correction of residuals allowed(**e6&e9; e9& e12**)	18.7(10; 0.044)	0.991	0.980	0.029(0.005,0.049)

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Table 5.2: Parameter estimates of unconditional LGM for the gain in length of infants in Butajira between the age of two and 18 months

Parameters of interest	*Estimate*	*Standard error*	*p-value*
Mean			
Intercept	57.30	0.141	<0.001
Slope	17.52	0.171	<0.001
Regression coefficients			
4 month time point	0.213	0.017	<0.001
6 month time point	0.403	0.007	<0.001
9 month time point	0.562	0.007	<0.001
12 month time point	0.724	0.006	<0.001
Covariance			
Slope and intercept	-0.123	0.822	0.881
Measurement error at			
6 and 9 months	1.214	0.382	0.001
9 and 12 months	1.757	0.378	<0.001
Variances			
Intercept	2.977	0.670	<0.001
Slope	7.070	1.402	<0.001
Measurement error			
2 month	12.013	0.905	<0.001
4 months	12.876	1.486	<0.001
6 month	8.538	0.507	<0.001
9 month	9.586	0.551	<0.001
12 month	7.756	0.463	<0.001
18 month	4.318	0.527	<0.001

An average infant in this population attains a mean length of 57.3 cm at two months of age, gains 17.5cm over the following 16 months and gains significant proportion between two months and any one of the subsequent time points of the total gain in length (see regression coefficients in table 5.2). There is significant variability both in initial length of infants at the age of two months and the overall gain in length between two and eighteen months of age (p = 0.001). However, there is no significant association between the initial length and the overall length that an average infant gains within the following 16 months. Of the total increase in length of an average infant 21.3% occurs between two and four months of age, 19.0% (i.e. 0.403 – 0.213 = 0.19) occurs between four and six months of age, 15.9% (i.e. 0.562 – 0.403 = 0.159) occurs between six and nine months of age, 16.2% (i.e. 0.724 – 0.562 = 16.2) occurs between nine and twelve months of age, and 27.6% (i.e. 1 – 0.724 = 0.276) occurs in the last six months (table 5.2). At each infant's age considered in this analysis there is significant variability in the attained length which is not accounted for by the LGM parameters and this is relatively higher at the ages of two and four months of age (see estimates of measurement error in table 5.2).

5.2.2 Weight gain of infants

Summary of fit indices obtained after fitting different LGMs to the weight of infants between the ages of two and eighteen months of age are summarized in table 5.3. If we base our model selection criteria only on the values of CFI and TLI quadratic and non-linear LGMs fit to the data fairly well. Considering the values of all four indices of global fit and allowing residuals to be serially correlated a non-linear LGM which is not quadratic gives a better description of the weight data ($\chi^2(df = 10) = 12.5$, p=0.225; TLI = 0.996, CFI = 0.996, RMSEA = 0.015 ; 90%CI: 0.000 to 0.039).

Table 5.3: Values of fit indices for selected Latent Growth Models (LGM) for the Weight gain of infants

Model description	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA(90%CI)***
Linear growth with Independence residuals	432.1(16; 0.001)	0.675	0.574	0.158(0.145,0.171)
Linear growth with Serial correlation of residuals allowed (**e6& e9**)	334.9(14; 0.001)	0.750	0.625	0.148(0.134,0.162)
Quadratic growth Independence of residuals	83.8(12; <0.001)	0.944	0.902	0.076(0.061,0.091)
Quadratic growth Serial correlation of residuals allowed (**e6, e9; e12**)	58.3(10; <0.001)	0.962	0.921	0.068(0.052,0.085)
Non-linear growth Independence of residuals allowed	59.5(12; <0.001)	0.963	0.935	0.062(0.043,0.078)
Non-linear growth Serial correction of residuals allowed(**e6,e9,e12**)	12.5(10; 0.225)	0.998	0.996	0.015(0.000,0.039)

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Parameter estimates of unconditional non-linear LGM fitted to the weight of infants is summarized in table 5.4. A two month old average infant weighed 5.13kg with significant variability between weights of two month old infants (P <0.001). The overall weight gain of an average infant from two to eighteen months of age amounts to 4.49kg and this weight gain varies significantly across infants (P<0.001). For an average infant there is no significant covariance between the initial weight and the overall weight gain during the study period. There is significant variability in the attained individual weight at each age studied which is not accounted for by the latent growth trajectory parameters and this variability generally increased as infants get older. The finding also shows that 20.8% of the total weight change occurred between two and four months of age and 28.5% of the weight change was between the ages of 12 and 18 months. Of the total mean weight change that occurred from two to 18 months of age the age ranges of 4 to 6 months, 6 to 9 months and 9 to 12 months account for 17.1%, 17.0% and 17.3%, respectively.

Table 5.4: Parameter estimates of non-linear unconditional LGM for the gain in weight of infants

Parameters of interest	*Estimate*	*Standard error*	*p-value*
Mean			
Intercept	5.13	0.020	<0.001
Slope	4.49	0.048	<0.001
Regression coefficients			
4 month time point	0.218	0.012	<0.001
6 month time point	0.372	0.007	<0.001
9 month time point	0.542	0.009	<0.001
12 month time point	0.715	0.008	<0.001
Covariance			
Slope and intercept	0.075	0.052	0.145
Measurement error at			
6 and 9 months	0.184	0.032	<0.001
9 and 12 months	0.177	0.037	<0.001
Variances			
Intercept	0.309	0.038	<0.001
Slope	0.769	0.106	<0.001
Measurement error			
2 month	0.319	0.905	<0.001
4 months	0.387	1.486	<0.001
6 month	0.567	0.507	<0.001
9 month	0.975	0.551	<0.001
12 month	0.757	0.463	<0.001
18 month	0.871	0.527	<0.001

5.2.3. Length-for-age z score and weight-for-age z score

Summary of fit indices for three unconditional LGMs (i.e. linear, quadratic and non-linear which is not quadratic) fitted to length-for-age z score and weight-for-age z score are summarized in tables 5.5 and 5.6, respectively. Quadratic LGM with independent residuals fits the data better than the other two models. When the assumption of independent residuals is relaxed and serial correlation is allowed the model fit was significantly improved both for length-for-age z score (Table 5.5) ($\chi^2(df = 10) = 10.7$, p<0.000; TLI = 0.999, CFI = 0.999, RMSEA = 0.008 ; 90%CI: 0.000 to 0.035) and weight-for-age z score (Table 5.6)($\chi^2(df = 11) = 64.4$, p<0.000; TLI = 0.929, CFI = 0.963, RMSEA = 0.069 with 90%CI: 0.050 to 0.082). Linear LGM resulted in poor fit

but that of non-linear LGM resulted in a comparable fit to the quadratic model which is a more parsimonious model.

Table 5.5: Fit indices for linear, quadratic and non-linear LGM in modelling length-for-age z score of infants between 2 and 18 months of age in Butajira

Model description	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA(90%CI)***
Linear with independent residuals	233.6(16; <0.001)	0.785	0.717	0.114(0.101,0.127)
Linear with serial correlation of residuals allowed (**e6, e9&e12)**	159.4(14; <0.001)	0.856	0.784	0.100(0.084,0.114)
Quadratic with independence of residuals	19.9(12; 0.069)	0.992	0.986	0.025(0.000,0.044)
Quadratic with serial correlation of residuals allowed (**e6, e9&e12**)	10.7(10; <0.000)	0.999	0.999	0.008(0.000,0.035)
Non-linear with independence of residuals allowed	42.2(11; <0.000)	0.969	0.941	0.052(0.036,0.069)
Non-linear with serial correction of residuals allowed (**e6, e9& e12**)	9.2(9; 0.416)	1.00	0.999	0.005(0.000,0.035)

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Table 5.6: Fit indices for linear, quadratic and non-linear LGM in modelling weight-for-age z score of infants between 2 and 18 months of age in Butajira

Model description	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA(90%CI)***
Linear with independent residuals	316.9(16; <0.000)	0.774	0.703	0.134(0.121,0.147)
Linear with serial correlation of residuals allowed	230.0(6; <0.000)	0.837	0.756	0.122(0.108,0.136)
Quadratic with independence of residuals	72.2(12; <0.001)	0.955	0.921	0.069(0.054,0.085)
Quadratic with serial correlation of residuals allowed (**e6&e9)**	60.4(11; <0.000)	0.963	0.929	0.066(0.050,0.082)
Non-linear with independence of residuals allowed	77.8(12; <0.000)	0.951	0.913	0.072(0.058,0.088)
Non-linear with serial correction of residuals allowed (**e6, e9 & e12**)	61.8(10; 0.000)	0.961	0.918	0.070(0.054,0.088)

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Estimates of model parameters obtained from best fitting unconditional quadratic LGM for length-for-age z score and weight-for-age z score are summarized in table 5.7. Compared to the 2006 WHO reference population an average infant in the P-MaMiE study attained significantly lower mean height-for-age z score (α = -0.239; SE = 0.051) and mean weight-for-age z score (α = -0.552; SE = 0.038). There is significant inter individual variability in both growth measurements. Length-for-age z score and weight-for-age z score of an average infant continue to fall as the age progresses with significant variability between infants in cases of weight-for-age z score and without significant variability between infants in case of length-for-age z score. An initial mean weight-for-age z score of an average infant is inversely related with the rate at which this score falls. During the first eighteen months of infancy there is significant variability in height-for-age z score and weight-for-age z score which is not accounted for by the best fitting quadratic LGM. If we use percentage of variability explained by the LGM to rank the fit of models, the quadratic LGM fitted better to weight-for-age z score than to length-for-age z score and, generally accounted for more variability of both growth measures as infants become older than at an early ages of infancy.

Table 5.7: Parameter estimates obtained from unconditional quadratic LGM of length-for-age z score and weight-for-age z score without serially correlated measurement errors

Description of Estimated Model Parameters	*Length-for-age z score*			*Weight-for-age z score*		
	Parameter values	SE	P-value	Parameter values	SE	P-value
(a) Means						
Intercept	-0.239	0.051	0.000	-0.552	0.038	0.000
Linear slope	-0.265	0.013	0.000	-0.131	0.009	0.000
Quadratic slope	0.008	0.001	0.000	0.007	0.001	0.000
(b) Variances						
Intercept	0.500	0.188	0.008	0.806	0.100	0.000
Linear slope	0.011	0.013	0.363	0.027	0.005	0.000
Quadratic slope	0.000	0.000	0.648	0.000	0.000	0.000
Measurement errors at						
2 months	2.169	0.219	0.000	0.489	0.095	0.000
4 months	0.813	0.125	0.000	0.585	0.089	0.000
6 months	1.763	0.108	0.000	0.858	0.058	0.000
9 months	1.424	0.115	0.000	1.005	0.065	0.000
12 months	1.191	0.115	0.000	0.669	0.053	0.000
18 months	0.567	0.261	0.030	0.539	0.184	0.003
(c) Covariance						
Intercept & linear slope	0.021	0.041	0.611	-0.051	0.021	0.014
intercept & Quadratic slope	-0.001	0.002	0.525	0.002	0.001	0.041
Linear slope & Quadratic slope	0.000	0.001	0.527	-0.001	0.000	0.000
Measurement errors at						
6 and 9 months	0.184	0.080	0.021	0.174	0.050	0.000
9 and 12 months	0.252	0.095	0.008	-	-	
(d) R^2 (coefficient of determination) at						
2 months	0.187			0.622		
4 months	0.430			0.548		
6 months	0.302			0.468		
9 months	0.411			0.474		
12 months	0.494			0.604		
18 months	0.680			0.600		

SE = Standard error

5.2.4. Stunting and underweight

Six unconditional LGMs were fitted using Mplus to describe the change in stunting and underweight of infants from two to eighteen months of age and goodness of fit indices of these models are summarized in table 5.8. Except chi-square which is well known to be influenced by large sample size the other three fit indices are within the recommended range for the best fitting model (CFI and TLI above 0.95 and RMSEA <0.06). When models are ranked using RMSEA quadratic and non-linear LGMs come first and second for underweight. For stunting non-linear LGM becomes first and quadratic fit becomes the second.

Table 5.8: Fit indices for linear, quadratic and non-linear Latent Growth Models in modelling stunting of infants between two and 18 months of age in Butajira

Growth measure	***Model description***	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA***
Stunting					
	Linear	22.11(9; 0.009)	0.974	0.965	0.039
	Quadratic	12.74(6; 0.047)	0.986	0.973	0.034
	Non-linear	13.62(7; 0.058)	0.987	0.977	0.031
Underweight					
	Linear	29.92(10;0.001)	0.977	0.975	0.046
	Quadratic	6.56(6; 0.364)	0.999	0.999	0.010
	Non-linear	12.63(7; 0.082)	0.994	0.990	0.029

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Proportion of variability explained by the underlying latent growth variable (i.e. r^2) is summarized in Table 5.9 and figure 5.1. Based on these values model fit was generally better as infants get older both for stunting and underweight. The three LGMs described underweight in better ways than stunting. Starting at nine months of age the models have an overall ability to describe more than 50% of the variability within each outcome variable.

Table 5.9: Proportion of variability in stunting and underweight which is explained by the underlying LGM

Growth outcome	*Age of infants in months*	*Linear*	*Quadratic*	*Non-linear*
Stunting				
	Two months	0.077	0.672	0.093
	Four months	0.217	0.441	0.045
	Six months	0.252	0.277	0.233
	Nine months	0.514	0.556	0.564
	Twelve months	0.624	0.630	0.619
	Eighteen months	0.457	0.585	0.438
Underweight	Two months	--	0.626	0.707
	Four months	0.549	0.690	0.546
	Six months	0.524	0.789	0.544
	Nine months	0.481	0.553	0.536
	Twelve months	0.687	0.714	0.758
	Eighteen months	0.739	0.694	0.637

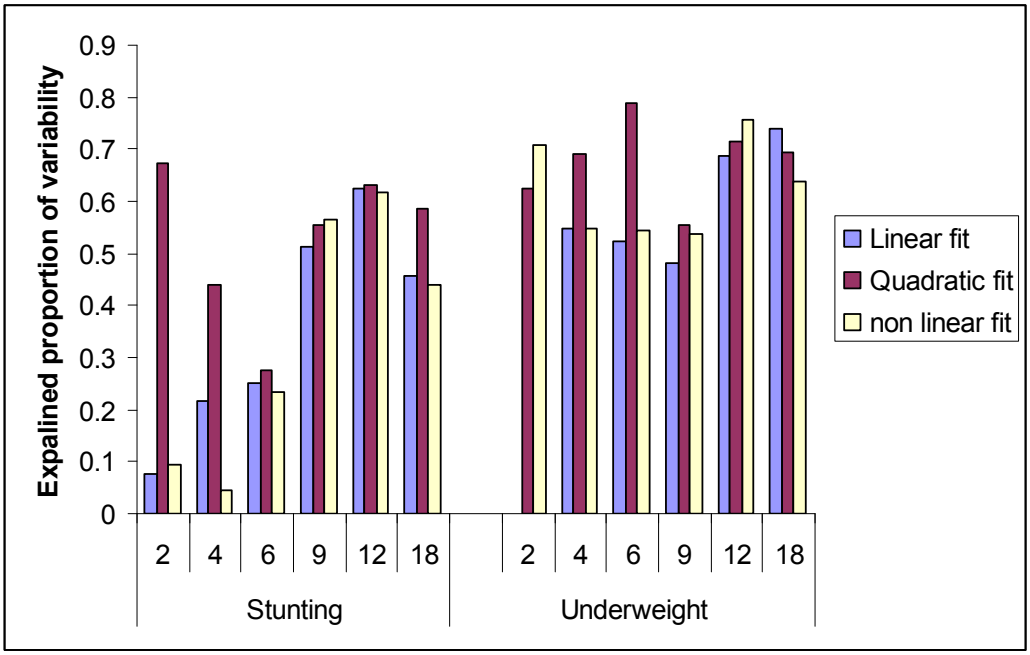

Figure 5.1: Proportion of variability in the infant undernutrition explained by the underlying latent growth variable

Three different LGMs resulted in similar likelihood of stunting as well as underweight for a two month old average infant which are all significantly greater than zero (see the magnitude of thresholds in table 5.10 and model predicted probabilities in figure 5.2). In relative terms the probability of being stunted is higher than the probability of being underweight. Variability of model predicted probabilities of stunting/underweight between two month old infants is model dependent. There is significant variability between infants in the probability of underweight predicted from quadratic and non-linear LGMs and in the probability of stunting predicted from quadratic LGM. However, there is homogeneity in their probability of stunting/underweight predicted from linear LGM.

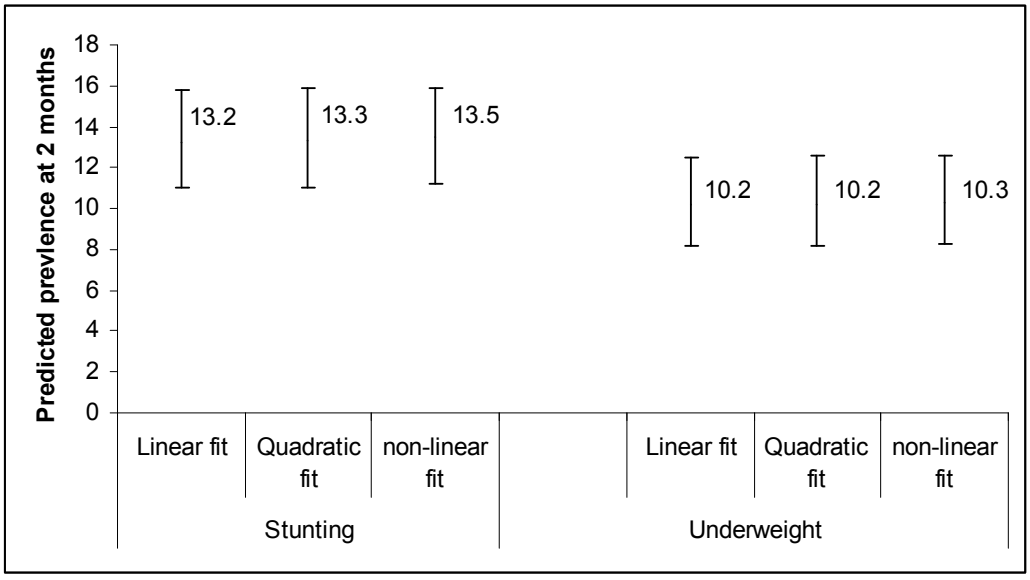

Figure 5.2: Model predicted probabilities and their 95% confidence interval of a two month average infant being undernourished in Butajira, Ethiopia

Table 5.10: Parameter estimates from three different unconditional LGMs for stunting and underweight of infants in Butajira, Ethiopia

Growth outcome	*Description of Estimated Model Parameters*	*Linear change model*	*Quadratic change model*	*Non-linear change model*
		Estimate (95% CI)	Estimate (95% CI)	Estimate (95% CI)
Stunting	**(a) Means**			
	Threshold at 2 months	1.115(1.001, 1.229)	1.111(0.997, 1.224)	1.103(0.989, 1.217)
	Linear term	0.113(0.099, 0.127)	0.113(0.099, 0.128)	1.428(1.295, 1.561)
	Quadratic term		0.00(0.000, 0.000)	
	(b) Variances			
	Intercept	0.077(-0.109, 0.263)	0.672(0.235, 1.110)	0.093(-0.106, 0.292)
	Linear term	0.006(0.001, 0.011)	0.050(0.020, 0.080)	0.146(-0.219, 0.512)
	Quadratic term		0.00(0.000, 0.000)	
	(C) Covariance			
	Linear term with Intercept	0.018(-0.003, 0.038)	-0.142(-0.244, -0.039)	0.100(-0.138, 0.338)
	Linear term with quadratic term		-0.003(-0.004, -0.001)	
	Intercept with Quadratic term		0.009(0.003, 0.014)	
	(D) Regression weights of linear slope			
	Two months	0	0	0
	four months	2	2	-0.157(-0.821, 0.508)
	Six months	4	4	0.386(0.282, 0.490)
	Nine months	7	7	0.659(0.603, 0.715)
	12 months	10	10	0.749(0.696, 0.802)
	18 months	16	16	1
Underweight				
	(a) Means			
	Threshold at 2 months	1.270(1.149,1.392)	1.268(1.147,1.389)	1.267(1.146,1.388)
	Linear term	-0.730(-1.830,0.370)	0.466(0.265,0.668)	0.382(0.245,0.520)
	Quadratic term		-0.069(-0.108,-0.030)	
	(b) Variances			
	Intercept	1.059(-0.235,2.353)	0.626(0.018,1.233)	0.707(0.129,1.285)
	Linear term	0.674(-1.115,2.462)	0.096(-0.017,0.209)	0.442(-0.090,0.974)
	Quadratic term		0.003(-0.001,0.007)	
	(C) Covariance			
	Linear term with Intercept	0.303(-0.137,0.743)	-0.215(-0.444,0.014)	-0.256(-0.787,0.275)
	Linear term with quadratic term		-0.015(-0.034,0.004)	
	Intercept with Quadratic term		0.032(0.006,0.059)	
	(D) Regression weights of linear slope			
	Two months	0	0	0
	four months	2	2	0.236(-0.394,0.865)
	Six months	4	4	1.061(0.843,1.279)
	Nine months	7	7	1.244(0.983, 1.505)
	12 months	10	10	1.247(1.018, 1.476)
	18 months	16	16	1

Results from quadratic LGM show significant increase in the likelihood of stunting of an average infant from two to eighteen months of age with significant variability between infants, significant negative correlation with initial value at the age of two months and positive correlation with quadratic term. Similarly, there is significant increase in the likelihood of underweight over time without significant variability between infants. Linear LGM resulted in statistically significant increase in the likelihood of stunting which varies significantly between infants but the increase in the likelihood of underweight as well as the variability in the likelihood of underweight between infants are not statistically significant. Under the non-linear LGM fit there is a significant age specific linear slope which is homogenous between two month old infants. The likelihood of stunting of an average infant continues to increase from two to eighteen months of age but the likelihood of underweight is decreased between the ages of twelve and eighteen months of age.

5.3 Summary of the findings

To determine the best fitting unconditional LGMs for infant anthropometric data we modelled continuous growth outcomes (i.e. weight, length, weight-for-age and length-for-age) using AMOS version 7 (Arbuckle 1995-2006) and binary growth outcomes (i.e. stunting and underweight) using Mplus version 5 (Muthen and Muthen 1998-2009). For each growth outcome we investigated three candidate LGMs (i.e. linear, quadratic and non-linear which is not quadratic). Selection of the best fitting LGM from the three candidate growth models was based on comparison of models with the assumption of independence of residuals. Further improvement of the model fit was investigated by relaxing the independence of residual assumptions. Goodness of model fit was assessed using chi-square, TLI, CFI, and RMSEA. Of the three candidate LGMs models quadratic fit was selected to describe weight-for-age and length-for-age, and non-linear LGM which is not quadratic was selected to describe length and weight. With the assumption of normal distribution of the underlying latent variables probit scale was used to model change in the probabilities of stunting and underweight and non-linear LGM which is not quardartic was selected to describe the two outcomes.

A two month old average infant in this population (a) attains a weight of 5.13kg and length of 57.3cm, (b) scores 0.24 and 0.55 standard deviation units below the median score of WHO reference population on its length-for-age z score and weight-for-age z score, respectively, and (c) has a 10.3% probability of being underweight and a 13.5% probability of being stunted. There is significant variability between two month old individual infants in their attained weight, length, weight-for-age z score, length-for-age z score and probability of stunting.

Between two and eighteen months of age the average infant gains 4.5kg in weight and 17.5cm in length. This increase in weight and length significantly vary between infants. Compared to the WHO reference standards an average infant in this population attains smaller length-for-age z score and weight-for-age z score during the whole study. This negative change in length-for-age z is homogeneous between infants but there is significant variability in the case of weight-for-age z. At eighteen months of age probability of stunting of an average infant increased to 62.7%. Probability of underweight reached 21.5% at one year and goes down to 18.8% at 18 months of age. There is no significant variability between infants in the rate at which the probability of stunting and underweight changes between two and eighteen months of age.

CHAPTER 6: RESULTS - CONDITIONAL LGMs

6.1 Introduction

Results from several LGMs fitted to each of the six infant growth outcomes (i.e. length, weight, length-for-age z, weight-for-age z, stunting and underweight) are summarized in this chapter. As in chapter 4 the exposure of main interest was perinatal CMD. Unlike the previous chapter perinatal CMD was treated as two separate variables, namely, antenatal CMD and postnatal CMD rather than three variables. To take account of possible continuity of CMD from antenatal to postnatal period (i.e. to evaluate the effect of persistent CMD) we deviated from what we have done in chapter 4 and the flexibility of LGM technique was exploited. Postnatal CMD was allowed to have a mediation role from antenatal CMD to infant growth on top of its independent effect on infant growth.

The overall result is summarized under the following three headings:

- Partially adjusted effects of maternal CMD on LGM parameters:- the focus of this section is on whether maternal CMD had any significant effect on initial infant growth and the rate at which growth of infants change over time without taking account of the role of any other risk factors. Direct effects of antenatal, postnatal and persistent CMD on the LGM parameters of infant growth are summarized in this section.
- Mediating effects of maternal and selected infant characteristics:- this section is a continuation of the previous section and it focuses on direct effects and mediating roles of birth weight, history of infant illnesses before two months follow-up, and early infant feeding practices of the mother. Within this context direct and indirect effects of perinatal CMD were investigated and summarized in this section.
- Fully adjusted effects of perinatal CMD and other risk factors:- this section has summary results from the final LGMs for each outcome. Direct and indirect effects of CMD and direct effects of pre-specified risk factors on LGM parameters are summarized here.

6.2 Partially adjusted effects of maternal CMD on LGM parameters

6.2.1 Goodness of fit of the LGMs

Based on the findings of the unconditional LGMs two sets of conditional LGMs were considered to investigate the effects of perinatal CMD on infant growth:

(a) quadratic LGM for weight-for-age and length-for-age z-scores and,

(b) non-linear LGM which is not quadratic for weight, length, underweight (i.e. probit of underweight) and stunting (i.e. probit of stunting).

SRQ-20 was treated as a binary variable and as a continuous variable in assessing the effect of CMD on each of the six anthropometric growth outcomes. This has resulted in two LGMs for each growth outcome. Table 6.1a has summarized the values of fit indices obtained after fitting twelve conditional LGMs. In almost all models chi-square was significant as is usually found for this sort of model but the other three fit indices were within the recommended range for the best fitting model.

Table 6.1a Fit indices for conditional LGMs of infants in Butajira, Ethiopia, where the only exposure measures are antenatal and postnatal CMD

Growth measure	***Measure of CMD***	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA***
Length					
	SRQ as binary	29.59(19;0.057)	0.990	0.981	0.023
	SRQ score	40.37(19;0.003)	0.980	0.963	0.033
Weight					
	SRQ as binary	20.41(19;0.370)	0.999	0.998	0.008
	SRQ score	30.82(19;0.042)	0.992	0.984	0.024
Length-for-age					
	SRQ as binary	22.21(16;0.136)	0.994	0.988	0.019
	SRQ score	23.59(16;0.099)	0.994	0.986	0.021
Weight-for-age					
	SRQ as binary	56.46(16;0.000)	0.972	0.936	0.049
	SRQ score	58.56(16;0.000)	0.971	0.935	0.050
Stunting					
	SRQ as binary	20.67(14;0.111)	0.988	0.981	0.021
	SRQ score	33.42(14; 0.003)	0.977	0.963	0.037
Underweight					
	SRQ as binary	15.61(14;0.338)	0.998	0.997	0.001
	SRQ score	20.09(13;0.093)	0.994	0.991	0.023

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

When perinatal CMD was included in the models without mediating or confounding variables it explained a small proportion of variability of the LGM parameters of each growth outcome regardless of whether SRQ-20 was used as a continuous or as a binary variable (table 6.1b). The exception was that of length-for-age where 30.3% of the variability of the linear slope term was explained when SRQ-20 was treated as continuous variable where as only 9.4% of the corresponding variability was explained when SRQ-20 was treated as a binary variable. In relative terms the fit of all LGMs was better as infants get older and infant weight (standardized or unstandardized) was described by these models better than infant length (standardized or unstandardized)

Table 6.1b: Proportion of total variation in the LGM parameters and measured growth outcomes explained by perinatal CMD

Measured outcomes	*Model parameters*			*Months of growth measurement*					
	I	S	Q	2	4	6	9	12	18
	SRQ-20 was treated as a continuous exposure variable								
Length	0.014	0.008		0.197	0.199	0.320	0.345	0.456	0.691
Weight	0.012	0.015		0.557	0.512	0.453	0.381	0.509	0.587
length-for-age	0.089	0.303	0.601	0.197	0.426	0.302	0.408	0.492	0.670
Weight-for-age	0.015	0.042	0.037	0.579	0.547	0.453	0.417	0.540	0.740
Stunting	0.091	0.073		0.087	0.043	0.232	0.567	0.622	0.432
underweight	0.025	0.03		0.757	0.551	0.545	0.516	0.790	0.633
	SRQ-20 was treated as a binary exposure variable								
Length	0.005	0.002		0.194	0.199	0.319	0.346	0.457	0.693
Weight	0.011	0.007		0.554	0.513	0.453	0.381	0.509	0.587
length-for-age	0.032	0.094	0.194	0.194	0.431	0.302	0.410	0.493	0.672
Weight-for-age	0.007	0.019	0.015	0.573	0.548	0.454	0.417	0.538	0.746
Stunting	0.019	0.020		0.087	0.046	0.236	0.567	0.614	0.439
underweight	0.076	0.048		0.730	0.545	0.552	0.529	0.763	0.636

Note: I = Intercept term; S = slope term; Q = Quadratic term

6.2.2 Direct and indirect effects of perinatal CMD on growth of infants

Estimates of direct effects of perinatal CMD on LGM parameters of infant growth from twelve partially adjusted conditional LGMs are summarized in Table 6.2a. There were no significant direct effects of antenatal and postnatal CMD on infant length or underweight. The effects on the other four growth outcomes were slightly influenced by the scale of SRQ-20 in the LGM.

When SRQ-20 was treated as a dimensional measure of CMD (i.e. SRQ-20 as a continuous exposure variable),
(a) antenatal CMD had significant

- negative effect on initial value of the probit of stunting and on the overall change of weight between two and eighteen months of age of an average infant,
- positive effect on the overall change of the probit of stunting between two and eighteen months of an average infant**,**
- positive effect on the intercept and quadratic terms and significant negative effect on the linear slope of both length-for-age and weight-for-age z of an average infant,

(b) postnatal CMD had significant negative effect on initial length-for-age z and quadratic term, and significant positive effect on linear slope term of length-for-age z of an average infant.

When SRQ-20 was dichotomized and included as binary exposure variable within the LGM,

- antenatal CMD had significant negative effect on linear slope term and significant positive effect on quadratic term of length-for-age z of an average infant and
- postnatal CMD had significant negative effect on initial value and quadratic term, and had significant positive effect on linear slope term of length-for-age z of an average infant.

There was significant continuity of maternal CMD from pregnancy to postnatal period and significant effect of persistent CMD on infant growth was only on length-for-age z

score (Table 6.2b). When SRQ-20 was used as a dimensional measure of CMD, persistent CMD had significant negative effect on initial term and quadratic term, and positively associated with linear slope term of length-for-age z of an average infant. When SRQ-20 was considered as a binary exposure variable the effects of persistent CMD on initial value and instantinuous rate of change of length-for-age z at two months were marginally non-significant.

Table 6.2a: Unadjusted effects of antenatal and postnatal CMD on growth trajectories of infants in Butajira, Ethiopia

Predictor variable	Intercept factor			*Slope factor*			*Quadratic factor*			*Postnatal CMD*		
Outcome	Estimate	SE`	p-value	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
Length												
Antenatal CMD (SRQ binary)	0.304	0.363	0.402	-0.329	0.508	0.517				0.219	0.021	0.000
Postnatal CMD (SRQ binary)	-0.507	0.560	0.365	0.186	0.783	0.812						
Antenatal CMD (SRQ score)	0.066	0.038	0.082	-0.079	0.054	0.141				0.294	0.023	0.000
Postnatal CMD (SRQ score)	-0.054	0.052	0.300	0.054	0.073	0.457						
Weight												
Antenatal CMD (SRQ binary)	0.062	0.085	0.466	-0.255	0.147	0.083				0.219	0.021	0.000
Postnatal CMD (SRQ binary)	0.229	0.131	0.081	0.055	0.226	0.806						
Antenatal CMD (SRQ score)	0.013	0.009	0.142	-0.038	0.016	0.015				0.293	0.023	0.00
Postnatal CMD (SRQ score)	0.015	0.012	0.211	0.007	0.021	0.744						
Length-for-age												
Antenatal CMD (SRQ binary)	0.321	0.167	0.055	-0.091	0.041	0.030	0.004	0.002	0.042	0.219	0.021	0.000
Postnatal CMD (SRQ binary)	-0.509	0.259	0.049	0.120	0.064	0.062	-0.007	0.003	0.036			
Antenatal CMD (SRQ score)	0.063	0.017	0.000	-0.017	0.004	0.000	0.001	0.000	0.000	0.294	0.023	0.000
Postnatal CMD (SRQ score)	-0.074	0.024	0.002	0.022	0.006	0.000	-0.001	0.000	0.000			
Weight-for-age												
Antenatal CMD (SRQ binary)	0.117	0.125	0.349	-0.057	0.030	0.062	0.002	0.002	0.144	0.219	0.021	0.000
Postnatal CMD (SRQ binary)	0.227	0.192	0.238	0.012	0.047	0.796	-0.001	0.003	0.629			
Antenatal CMD (SRQ score)	0.030	0.013	0.024	-0.009	0.003	0.005	0.000	0.000	0.030	0.293	0.023	0.000
Postnatal CMD (SRQ score)	0.011	0.018	0.529	0.005	0.004	0.237	0.000	0.000	0.134			
Stunting												
Antenatal CMD (SRQ binary)	-0.138	0.198	0.486	0.197	0.304	0.517				1.220	0.150	0.000
Postnatal CMD (SRQ binary)	0.024	0.114	0.831	0.031	0.164	0.848						
Antenatal CMD (SRQ score)	-0.032	0.015	0.030	0.053	0.024	0.024				0.293	0.014	0.000
Postnatal CMD (SRQ score)	0.013	0.021	0.550	-0.027	0.029	0.356						
Underweight												
Antenatal CMD (SRQ binary)	0.103	0.624	0.870	0.295	0.594	0.621				1.220	0.150	0.000
Postnatal CMD (SRQ binary)	-0.436	0.457	0.659	0.217	0.445	0.374						
Antenatal CMD (SRQ score)	-0.051	0.055	0.352	0.067	0.054	0.216				0.293	0.014	0.000
Postnatal CMD (SRQ score)	-0.073	0.094	0.434	0.038	0.086	0.658						

SE = Standard error

Table 6.2b**:** The effect of persistent CMD on growth trajectories of infants in Butajira, Ethiopia

Predictor variable	**Intercept factor**			***Slope factor***			*Quadratic factor*		
Outcome	Test statistic	SE`	p-value	Test statistic	SE	p-value	Test statistic	SE	p-value
Length									
SRQ as a binary	-0.902	0.123	0.367	-0.237	0.172	0.812			
SRQ as a score	-1.035	0.015	0.301	0.738	0.021	0.460			
Weight									
SRQ as a binary	1.724	0.029	0.084	0.243	0.050	0.808			
SRQ as a score	1.244	0.004	0.213	0.333	0.006	0.739			
Length-for-age									
SRQ as a binary	-1.931	0.058	0.054	1.845	0.014	0.065	-2.277	0.000	0.023
SRQ as a score	-2.997	0.007	0.003	3.524	0.002	0.000	-12.739	0.000	0.000
Weight-for-age									
SRQ as a binary	1.175	0.043	0.240	0.255	0.010	0.799	-0.333	0.000	0.739
SRQ as a score	0.610	0.005	0.542	1.244	0.001	0.213	-1.488	0.000	0.239
Stunting									
SRQ as a binary	0.210	0.139	0.833	0.189	0.200	0.850			
SRQ as a score	0.619	0.006	0.536	-0.930	0.009	0.352			
Underweight									
SRQ as a binary	-0.948	0.561	0.343	0.487	0.544	0.626			
SRQ as a score	-0.776	0.028	0.438	0.442	0.025	0.659			

SE = Standard error

6.3 Mediating effects of early infant feeding practices, infant illness and birth weight in modelling the influence of CMD on infant growth

6.3.1 Goodness of fit of the mediation models

Mediating roles of early infant feeding practices of the mother, infant illness and birth weight were investigated in modelling the effect of perinatal CMD on infant growth. In doing this twelve LGMs (two LGM for each growth outcome depending on which scale SRQ-20 was entered into the model) were fitted to the infant growth data and fit indices are presented in Table 6.3. All the LGMs provided good description of each growth outcome independent of how SRQ-20 was treated in the model. None of the models resulted in improper solutions of model parameters (i.e. no negative variance or correlation coefficient greater than one). Confidence intervals around point estimate of RMSEA for binary outcomes are not presented because Mplus does not provide them.

Table 6.3 Fit indices in evaluating mediating roles of selected maternal and infant characteristics in modelling the effect of CMD on infant growth in Butajira, Ethiopia

Growth measure	***Scale of SRQ***	***Chi-square (df; P-value)***	***CFI***	***TLI***	***RMSEA***
Length					
	Binary	90.4, (60; 0.007)	0.975	0.956	0.022(0.012,0.031)
	Continuous	105.9,(60, 0.000)	0.964	0.937	0.027(0.018,0.035)
Weight					
	Binary	68.9, (60; 0.202)	0.994	0.990	0.012(0.000,0.023)
	Continuous	86.4 ,(60, 0.014)	0.984	0.972	0.021(0.009,0.030)
Length-for-age					
	Binary	84.8, (53; 0.004)	0.975	0.951	0.024(0.014,0.033)
	Continuous	96.7,(53, 0.000)	0.968	0.937	0.028(0.019,0.037)
Weight-for-age					
	Binary	97.6, (52; 0.000)	0.972	0.943	0.029(0.020,0.038)
	Continuous	106.3, (52, 0.000)	0.968	0.936	0.032(0.023,0.040)
Stunting					
	Binary	75.3, (45; 0.000)	0.936	0.955	0.025
	Continuous	84.1, (45, 0.003)	0.938	0.956	0.029
Underweight					
	Binary	59.6., (45; 0.071)	0.981	0.986	0.018
	Continuous	64.0, (43, 0.020)	0.976	0.983	0.022

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

Proportion of variability in the LGM parameters and measured growth outcomes explained by the models are summarized in table 6.4. Maternal and infant characteristics included in the LGMs explained between 9.8% and 24.3% of the total variability of the initial growth status and between 3.0% and 37.7% of the total variability of the slope terms. For all the growth outcomes the proportion of explained variability was higher for initial values than for the slope terms except for stunting whose result was in the reverse order. Among the continuous growth outcomes a higher proportion of variability of the LGM parameters was explained in modelling weight or weight-for-age than in modelling length or length-for-age z scores. In case of the two binary growth outcomes the reverse was true with only one exception which appeared when SRQ-20 was treated as a binary variable.

As would be expected due to reduction in measurement errors the proportion of variability of measured growth outcomes explained by the model increased as infants get older with minimal effect from how SRQ-20 was treated in the model. The proportion of variability at each time point explained by the model was relatively higher for weight or its two derivatives compared to that of length and its two derivatives. The smallest proportion of variability explained by the model was for stunting at two and four months of age.

Table 6.4: Proportion of total variation in the LGM parameters and measured growth outcomes explained by the model

Measured outcomes	*Model parameters*			*Months of growth measurement*					
	I	S	Q	2	4	6	9	12	18
	SRQ-20 was treated as a continuous exposure variable								
Length	0.106	0.030	-	0.204	0.204	0.321	0.346	0.457	0.693
Weight	0.227	0.041	-	0.501	0.494	0.453	0.385	0.514	0.586
Length-for-age	0.168	0.151	0.162	0.286	0.475	0.332	0.484	0.563	0.770
Weight-for-age	0.226	0.159	0.155	0.582	0.547	0.451	0.416	0.540	0.739
Stunting	0.243	0.377	-	0.088	0.051	0.226	0.543	0.651	0.430
underweight	0.184	0.122	-	0.685	0.449	0.558	0.525	0.770	0.633
	SRQ-20 was treated as a binary exposure variable								
Length	0.098	0.022	-	0.202	0.204	0.320	0.346	0.458	0.696
Weight	0.222	0.033	-	0.502	0.496	0.453	0.385	0.514	0.587
Length-for-age	0.134	0.102	0.114	0.277	0.479	0.330	0.482	0.560	0.760
Weight-for-age	0.218	0.138	0.130	0.576	0.547	0.452	0.416	0.538	0.745
Stunting	0.186	0.298	-	0.087	0.048	0.227	0.545	0.641	0.436
underweight	0.200	0.137	-	0.699	0.526	0.562	0.533	0.743	0.642

Note: I = Intercept term; S = slope term; Q = Quadratic term

6.3.2 Direct effects of CMD and selected variables on infant growth

6.3.2.1 Length and weight

Direct effects of perinatal CMD, early infant feeding practices of mothers, early life infant illnesses and birth weight on the parameters of LGMs underlying infant length and weight are summarized in Tables 6.5 and 6.6, respectively. Antenatal CMD was significantly associated with a small weight gain of an average infant between two and eighteen months of age. However, it was not significantly associated with initial length or initial weight of an average infant and with overall gain in length of an average infant from two to eighteen months of age. Postnatal CMD was not significantly associated with any one of the parameters of LGMs underlying length and weight of an average infant. All these associations were independent of the scale used for SRQ-20.

Increased birth weight was significantly associated with higher initial length and heavier initial weight of an average infant. However, it was not significantly associated with overall change in weight or in length which had occurred between two and eighteen months of age. On average infants who have been reported receiving colostrum had shorter initial length and higher overall weight gain from two to eighteen months of age. Incidence of infant illnesses in the first two months of age, receiving pre-lacteal feed or delayed initiation of breast feeding did not have significant direct effects on parameters of LGMs of length and weight of an average infant. All these associations were true whether SRQ-20 was treated as a binary or as a continuous variable in the LGMs.

Table 6.5: Regression coefficients, standard errors and p-values from mediation LGM of infant length in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score						
Antenatal CMD (SRQ score)	0.068	0.038	0.077	-0.086	0.054	0.110
Postnatal CMD (SRQ score)	-0.046	0.053	0.382	0.039	0.074	0.604
Birth weight	1.048	0.353	0.003	0.740	0.517	0.152
Severe illness in the first 2 months	0.000	0.284	0.999	0.208	0.400	0.603
Diarrhoeal episode in the first 2 months	-0.430	0.264	0.104	-0.097	0.373	0.796
No pre-lacteal food	-0.176	0.545	0.747	-0.798	0.760	0.294
Colostrums given	-0.612	0.286	0.032	0.395	0.403	0.328
Breast feeding delayed for 1 hours	0.073	0.452	0.872	0.585	0.630	0.354
SRQ-20 used as a binary variable						
Antenatal CMD (SRQ binary)	0.343	0.364	0.346	-0.367	0.512	0.473
Postnatal CMD (SRQ binary)	-0.442	0.566	0.435	0.045	0.798	0.955
Birth weight	1.075	0.353	0.002	0.705	0.517	0.173
Severe illness in the first 2 months	0.028	0.283	0.923	0.174	0.399	0.662
Diarrhoeal episode in the first 2 months	-0.409	0.263	0.120	-0.103	0.372	0.782
Received pre-lacteal food	-0.193	0.548	0.724	-0.779	0.763	0.308
Colostrum given	-0.606	0.287	0.034	0.394	0.404	0.330
Breast feeding delayed for 1 hours	0.097	0.451	0.829	0.549	0.630	0.384

SE = Standard error

Table 6.6: Regression coefficients, standard errors and p-values from mediation LGM of infant weight in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score						
Antenatal CMD (SRQ score)	0.013	0.009	0.128	-0.040	0.015	0.008
Postnatal CMD (SRQ score)	0.018	0.012	0.139	0.005	0.021	0.797
Birth weight	0.635	0.074	0.000	0.016	0.139	0.907
Severe illness in the first 2 months	-0.039	0.063	0.541	0.133	0.114	0.243
Diarrhoeal episode in the first 2 months	-0.073	0.059	0.218	-0.122	0.106	0.251
No pre-lacteal food	0.012	0.122	0.922	-0.016	0.216	0.940
Colostrums given	0.018	0.063	0.780	0.280	0.114	0.014
Breast feeding delayed for 1 hours						
SRQ-20 used as a binary variable						
Antenatal CMD (SRQ binary)	0.073	0.082	0.377	-0.288	0.146	0.048
Postnatal CMD (SRQ binary)	0.227	0.126	0.072	0.052	0.226	0.81
Birth weight	0.636	0.074	0.000	0.010	0.139	0.941
Severe illness in the first 2 months	-0.029	0.063	0.649	0.118	0.113	0.299
Diarrhoeal episode in the first 2 months	-0.066	0.059	0.260	-0.137	0.106	0.194
Received pre-lacteal food	0.005	0.123	0.970	0.013	0.217	0.952
Colostrum given	0.014	0.063	0.824	0.284	0.115	0.013
Breast feeding delayed for 1 hours	-0.113	0.098	0.251	0.212	0.176	0.229

SE = Standard error

6.3.2.2 Length-for-age and weight-for-age

Summaries of direct effects, standard errors and p-values related to perinatal CMD, early infant feeding practices of the mother, infant illness during the first two months of age and birth weight on LGM parameters of standardized length and weight are provided in tables 6.7 and 6.8, respectively. Antenatal CMD was

(a) significantly associated with increased initial value and quadratic term of weight-for-age z independent of whether SRQ-20 was treated as a binary variable or as a continuous variable ,
(b) significantly associated with increased initial values and quadratic terms of length-for-age z when SRQ-20 was treated as continuous variable,
(c) inversely and significantly associated with linear terms of weight-for-age z and length-for-age z independent of how SRQ-20 was treated in the model.

Regardless of how SRQ-20 was treated in the model (binary or continuous) postnatal CMD was not significantly associated with the LGM parameters of weight-for-age z. However, it was inversely associated with initial value and quadratic term, and positively associated with increased linear slope of length-for-age z of an average infant.

Birth weight was positively associated with initial values and quadratic terms of both weight-for-age z and length-for-age z, but inversely associated with linear growth term of weight-for-age z independent of how SRQ-20 was treated in the model. The observed effect size of inverse association of birth weight and linear term of length-for-age z was not statistically significant. Infant diarrhoea in the first two months and delayed initiation of breast feeding were both significantly associated with linear slope and quadratic terms of length-for-age but their effects were in opposite directions. Diarrhoea had a negative effect on linear term and a positive effect on quadratic term but the direction of the effects of delayed initiation of breast feeding was the opposite. Denial of colostrums was associated with shorter initial length-for-age z. None of the early infant feeding practices of the mother were significantly associated with LGM parameters underlying weight-for-age z.

Table 6.7**:** Regression coefficients, standard errors and p-values from mediation LGM of infant length-for-age in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor			Quadratic factor		
Outcome	Estimate	SE`	p-value	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score									
Antenatal CMD (SRQ score)	0.066	0.018	0.000	-0.017	0.005	0.000	0.001	0.000	0.000
Postnatal CMD (SRQ score)	-0.076	0.025	0.002	0.023	0.006	0.000	-0.001	0.000	0.000
Birth weight	0.590	0.165	0.000	-0.069	0.043	0.107	0.004	0.002	0.048
Severe illness in the first 2 months	-0.156	0.131	0.234	0.031	0.034	0.359	-0.001	0.002	0.457
Diarrhoeal episode in the first 2 months	0.074	0.122	0.545	-0.084	0.031	0.007	0.005	0.002	0.006
No pre-lacteal food	-0.242	0.253	0.340	0.011	0.064	0.862	-0.002	0.003	0.646
Colostrums given	-0.266	0.132	0.043	0.033	0.034	0.331	-0.001	0.002	0.406
Breast feeding delayed for 1 hours	-0.271	0.209	0.195	0.127	0.053	0.015	-0.006	0.003	0.020
SRQ-20 used as a binary variable									
Antenatal CMD (SRQ binary)	0.382	0.169	0.024	-0.105	0.043	0.015	0.005	0.002	0.022
Postnatal CMD (SRQ binary)	-0.546	0.263	0.038	0.137	0.067	0.041	-0.008	0.004	0.021
Birth weight	0.604	0.166	0.000	-0.070	0.043	0.103	0.004	0.002	0.049
Severe illness in the first 2 months	-0.142	0.131	0.277	0.030	0.034	0.375	-0.001	0.002	0.453
Diarrhoeal episode in the first 2 months	0.070	0.122	0.566	-0.078	0.031	0.012	0.004	0.002	0.010
Received pre-lacteal food	-0.271	0.255	0.289	0.019	0.065	0.773	-0.002	0.003	0.575
Colostrum given	-0.262	0.132	0.048	0.032	0.034	0.349	-0.001	0.002	0.432
Breast feeding delayed for 1 hours	-0.278	0.210	0.185	0.133	0.053	0.012	-0.007	0.003	0.015

SE = Standard error

Table 6.8: Regression coefficients, standard errors and p-values from mediation LGM of infant weight-for-age z in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor			Quadratic factor		
Outcome	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score									
Antenatal CMD (SRQ score)	0.030	0.013	0.018	-0.009	0.003	0.004	0.000	0.000	0.034
Postnatal CMD (SRQ score)	0.015	0.018	0.405	0.006	0.004	0.190	0.000	0.000	0.095
Birth weight	0.979	0.110	0.000	-0.100	0.028	0.000	0.005	0.002	0.002
Severe illness in the first 2 months	-0.174	0.094	0.065	0.015	0.024	0.534	0.000	0.001	0.928
Diarrhoeal episode in the first 2 months	-0.077	0.087	0.381	-0.016	0.022	0.463	0.001	0.001	0.548
No pre-lacteal food	0.024	0.182	0.896	-0.033	0.046	0.470	0.002	0.003	0.403
Colostrums given	0.022	0.093	0.815	0.043	0.024	0.072	-0.002	0.001	0.182
Breast feeding delayed for 1 hours	-0.065	0.147	0.658	0.009	0.037	0.815	0.000	0.002	0.961
SRQ-20 used as a binary variable									
Antenatal CMD (SRQ binary)	0.137	0.122	0.261	-0.059	0.030	0.053	0.002	0.002	0.146
Postnatal CMD (SRQ binary)	0.223	0.187	0.231	0.016	0.047	0.728	-0.002	0.003	0.565
Birth weight	0.984	0.111	0.000	-0.101	0.028	0.000	0.005	0.002	0.002
Severe illness in the first 2 months	-0.154	0.094	0.102	0.013	0.024	0.574	0.000	0.001	0.918
Diarrhoeal episode in the first 2 months	-0.065	0.087	0.457	-0.015	0.022	0.483	0.001	0.001	0.625
Received pre-lacteal food	0.010	0.183	0.957	-0.027	0.046	0.550	0.002	0.003	0.456
Colostrum given	0.018	0.094	0.848	0.043	0.024	0.070	-0.002	0.001	0.183
Breast feeding delayed for 1 hours	-0.043	0.147	0.768	0.008	0.037	0.826	0.000	0.002	0.984

SE = Standard error

6.3.2.3 Stunting and underweight

Estimate of probits with their corresponding standard errors and p-values are summarized in Table 6.9 for stunting and in table 6.10 for underweight. When SRQ-20 was treated as a continuous exposure variable antenatal CMD was significantly associated with an increased probability of stunning at two months and with a larger overall change of the probit of stunting from two to eighteen months of age. An increase in birth weight was significantly associated with small overall change in the probit of stunting from two to eighteen months of age when SRQ-20 was treated as continuous variable but its effect was not statistically significant when SRQ-20 was treated as a binary variable. An inverse relationship between birth weight and probability of initial underweight was marginally non-significant. None of the other variables include in the model were significantly associated with the probability of underweight or stunting at the initial time or with the overall change in the probability of underweight or stunting from two to eighteen months of age.

Table 6.9**:** regression coefficients, standard errors and p-values for the predictors of change in probit of stunting in the mediation LGM in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score						
Antenatal CMD (SRQ score)	-0.031	0.016	0.049	0.060	0.025	0.015
Postnatal CMD (SRQ score)	0.001	0.023	0.982	-0.025	0.032	0.422
Birth weight	-0.088	0.130	0.497	-0.417	0.205	0.042
Severe illness in the first 2 months	0.052	0.073	0.470	-0.154	0.122	0.206
Diarrhoeal episode in the first 2 months	0.078	0.070	0.267	0.090	0.114	0.431
No pre-lacteal food	-0.037	0.122	0.761	0.176	0.187	0.346
Colostrums given	-0.005	0.077	0.948	0.024	0.114	0.836
Breast feeding delayed for 1 hours	0.001	0.100	0.989	-0.168	0.143	0.243
SRQ-20 used as a binary variable						
Antenatal CMD (SRQ binary)	-0.044	0.154	0.776	0.141	0.243	0.561
Postnatal CMD (SRQ binary)	-0.301	0.220	0.171	0.441	0.347	0.203
Birth weight	-0.119	0.130	0.359	-0.368	0.203	0.070
Severe illness in the first 2 months	0.042	0.074	0.564	-0.148	0.123	0.229
Diarrhoeal episode in the first 2 months	0.080	0.071	0.260	0.081	0.115	0.483
Received pre-lacteal food	-0.026	0.121	0.830	0.157	0.184	0.392
Colostrum given	-0.009	0.077	0.904	0.027	0.112	0.811
Breast feeding delayed for 1 hours	-0.014	0.101	0.890	-0.151	0.143	0.294

SE = Standard error

Table 6.10: regression coefficients, standard errors and p-values for the predictors the change in probit of underweight in the mediation conditional LGM in Butajira, Ethiopia

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
SRQ-20 used as a score						
Antenatal CMD (SRQ score)	-0.042	0.044	0.340	0.065	0.043	0.130
Postnatal CMD (SRQ score)	-0.078	0.088	0.380	0.025	0.078	0.745
Birth weight	-1.354	0.717	0.059	0.496	0.699	0.478
Severe illness in the first 2 months	0.163	0.191	0.394	-0.101	0.176	0.565
Diarrhoeal episode in the first 2 months	0.106	0.172	0.540	0.008	0.160	0.961
No pre-lacteal food	-0.014	0.286	0.962	0.048	0.248	0.847
Colostrums not given	0.054	0.164	0.742	-0.245	0.152	0.107
Breast feeding delayed for 1 hours	0.063	0.228	0.782	-0.169	0.206	0.411
SRQ-20 used as a binary variable						
Antenatal CMD (SRQ binary)	-0.088	0.409	0.830	0.405	0.403	0.315
Postnatal CMD (SRQ binary)	-1.720	1.165	0.140	1.042	1.156	0.367
Birth weight	-1.346	0.764	0.078	0.498	0.749	0.506
Severe illness in the first 2 months	0.142	0.195	0.466	-0.089	0.182	0.627
Diarrhoeal episode in the first 2 months	0.111	0.180	0.537	0.008	0.172	0.964
Received pre-lacteal food	0.002	0.286	0.995	0.025	0.251	0.922
Colostrum given	0.047	0.169	0.780	-0.249	0.159	0.117
Breast feeding delayed for 1 hours	0.069	0.230	0.797	-0.173	0.214	0.421

SE = Standard error

6.3.3 Direct and indirect effects of persistent CMD on infant growth

The effects of (a) antenatal CMD on early infant feeding practices, postnatal CMD and birth weight, (b) postnatal CMD on early infant illnesses, and (c) receiving colostrum or prelacteal feeding on delayed initiation of breast feeding are summarized in

- table 6.11 for continuous growth outcomes treating SRQ-20 as a continuous variable
- table 6.12 for continuous growth outcomes treating SRQ-20 as binary exposure variable
- table 6.13 for binary growth outcome variables (stunting and underweight) treating SRQ-20 as a continuous variable
- table 6.14 for binary growth outcomes treating SRQ-20 as a binary exposure

For continuous as well as binary growth outcome variables

- antenatal CMD was significant predictor of postnatal CMD and delayed initiation of breast feeding,
- postnatal CMD was significant predictor of infant diarrhoea and severe infant illness in the first two months of life, and
- providing colostrums or prelacteal feed to infants were significant predictors of delayed initiation of breast feeding.

All these associations hold whether SRQ-20 was treated as a continuous or as a binary exposure variable. When SRQ-20 was treated as a binary exposure variable antenatal CMD was a significant predictor of infants receiving prelacteal feeding.

Although birth weight was significantly associated with most of the growth outcome variables, the result presented in tables 6.11-6.14 showed that there was no significant effect of antenatal CMD on birth weight. Hence, birth weight was not a significant mediating variable for the associations of antenatal CMD and infant growth. Similarly, denial of colostrum and receiving prelacteal feeding did not mediate the associations of antenatal CMD and infant growth because each of these feeding practices were not significantly affected by antenatal CMD. Moreover, they were not assocated with LGM parameters of infant growth at all or their associations were only marginal.

Antenatal CMD had a significant positive effect on delayed initiation of breast feeding and persistent CMD had significant positive effect on early infant illnesses (i.e. diarrhoea and life threatening severe illness). These associations make delayed initiation of breastfeeding and other early infant feeding practices of the mother potential mediators for the associations of antenatal CMD and persistent CMD with infant growth, respectively. However, severe infant illness was not significantly associated with LGM parameters of any of the growth outcome variables. Infant diarrhoea and delayed initiation of breastfeeding had significant effects only on the linear slope and quadratic terms of length-for-age but acting in opposite directions. Diarrhoea had a positive effect on the linear slope and positive effect on quadratic term but delayed initiation of breast feeding had the opposite effects to these two terms. This result showed that persistent CMD had significant negative effect on linear slope term of LGM of length-for-age z of an average infant.

Table 6.11. Mediating role of selected maternal and child characteristics (continuous outcomes)

Mediating variables	*Parameter estimates*	*Predictor variables*			
		Antenatal CMD	**Postnatal**	**Colostrum**	**prelact**
SRQ-20 as a score					
Postnatal CMD (SRQ score)					
	Slope	0.293			
	SE	0.023			
	p-value	0.000			
Birth weight					
	Slope	0.004			
	SE	0.005			
	p-value	0.434			
Severe illness in the first 2 months					
	Slope		0.026		
	SE		0.006		
	p-value		0.000		
Diarrhoeal episode in the first 2 months					
	Slope		0.039		
	SE		0.006		
	p-value		0.000		
Received pre-lacteal food					
	Slope	0.001			
	SE	0.002			
	p-value	0.562			
Colostrums given					
	Slope	0.001			
	SE	0.004			
	p-value	0.729			
Breast feeding delayed for 1 hours					
	Slope	0.007		-0.075	0.126
	SE	0.003		0.021	0.038
	p-value	0.008		0.000	0.000

SE = Standard error

Table 6.12. Mediating role of selected maternal and child characteristics (continuous outcomes)

Mediating variables	*Parameter estimates*	*Predictor variables*			
		Antenatal CMD	**Postnatal**	**colostrum**	**Prelact**
SRQ-20 as a binary variable					
Postnatal CMD (SRQ score)					
	Slope	0.219			
	SE	0.021			
	p-value	0.000			
Birth weight					
	Slope	-0.002			
	SE	0.0048			
	p-value	0.971			
Severe illness in the first 2 months					
	Slope		0.187		
	SE		0.062		
	p-value		0.003		
Diarrhoeal episode in the first 2 months					
	Slope		0.347		
	SE		0.067		
	p-value		0.000		
Received pre-lacteal food					
	Slope	0.050			
	SE	0.021			
	p-value	0.020			
Colostrums given					
	Slope	0.028			
	SE	0.039			
	p-value	0.469			
Breast feeding delayed for 1 hours					
	Slope	0.065		-0.076	0.119
	SE	0.025		0.021	0.038
	p-value	0.009		0.000	0.002

SE = Standard error

Table 6.13 Mediating role of selected maternal and child characteristics (binary outcomes)

Mediating variables	*Parameter estimates*	*Predictor variables*			
		Antenatal CMD	**Postnatal**	**colostrum**	**prelact**
SRQ-20 as a score					
Postnatal CMD (SRQ score)					
	Slope	0.298			
	SE	0.015			
	p-value	0.000			
Birth weight					
	Slope	0.004			
	SE	0.006			
	p-value	0.442			
Severe illness in the first 2 months					
	Slope		0.079		
	SE		0.018		
	p-value		0.000		
Diarrhoeal episode in the first 2 months					
	Slope		0.106		
	SE		0.016		
	p-value		0.000		
Received pre-lacteal food					
	Slope	0.012			
	SE	0.022			
	p-value	0.575			
Colostrums given					
	Slope	0.005			
	SE	0.014			
	p-value	0.708			
Breast feeding delayed for 1 hours					
	Slope	0.043		-0.308	0.359
	SE	0.021		0.097	0.127
	p-value	0.036		0.001	0.005

SE = Standard error

Table 6.14. Mediating role of selected maternal and child characteristics (binary outcomes)

Mediating variables	*Parameter estimates*	*Predictor variables*			
		Antenatal CMD	**Postnatal**	**colostrum**	**prelact**
SRQ-20 as a binary variable					
Postnatal CMD (SRQ score)					
	Slope	0.223			
	SE	0.019			
	p-value	0.000			
Birth weight					
	Slope	0.002			
	SE	0.045			
	p-value	0.968			
Severe illness in the first 2 months					
	Slope		0.562		
	SE		0.202		
	p-value		0.005		
Diarrhoeal episode in the first 2 months					
	Slope		0.922		
	SE		0.192		
	p-value		0.000		
Received pre-lacteal food					
	Slope	0.391			
	SE	0.177			
	p-value	0.028			
Colostrums given					
	Slope	0.110			
	SE	0.145			
	p-value	0.447			
Breast feeding delayed for 1 hours					
	Slope	0.351		-0.313	0.343
	SE	0.190		0.097	0.124
	p-value	0.066		0.001	0.005

SE = Standard error

6.4. Fully adjusted effects of maternal CMD and other risk factors on the parameters of LGMs

6.4.1 Goodness of fit

Summary of four global fit indices obtained after fitting twelve fully adjusted conditional LGMs of infant growth is presented in table 6.15a. Whether SRQ-20 was entered into the models as a binary or as a continuous variable all models resulted in an acceptable description of the growth data. Not surprisingly, chi-square was statistically significant in most of the models. However, the other three fit indices were within the acceptable ranges for the best fitting LGM.

Table 6.15a: Fit indices for fully adjusted quadratic and non-linear conditional LGMs in modelling anthropometric growth of infants in Butajira, Ethiopia

Growth measure	*Scale of SRQ*	*Chi-square (df; P-value)*	*CFI*	*TLI*	*RMSEA*
Length					
	Binary	305.6, (185; 0.000)	0.935	0.903	0.025(0.020,0.030)
	Continuous	318.1, (187, 0.000)	0.932	0.900	0.026(0.021,0.031)
Weight					
	Binary	320.4, (204; 0.000)	0.950	0.927	0.023(0.018,0.028)
	Continuous	349.0, (205, 0.000)	0.940	0.913	0.026(0.021,0.031)
Length-for-age					
	Binary	319.3, (187; 0.000)	0.950	0.920	0.026(0.021,0.031)
	Continuous	334.0, (187, 0.000)	0.946	0.913	0.027(0.023,0.032)
Weight-for-age					
	Binary	257.8, (150; 0.000)	0.953	0.920	0.026(0.021,0.032)
	Continuous	268.5, (150, 0.000)	0.950	0.916	0.027(0.022,0.033)
Stunting					
	Binary	152.7, (119; 0.020)	0.933	0.916	0.018
	Continuous	168.6, (126, 0.007)	0.937	0.922	0.020
Underweight					
	Binary	138.5, (120; 0.119)	0.970	0.965	0.013
	Continuous	150.3, (121, 0.037)	0.962	0.956	0.017

CFI = Compartive fit index, TLI = Tuker-Lewis index, RMSEA = root mean squared error of approximation, df = degrees of freedom

The proportion of variability in LGM parameters and measured growth outcomes explained by the fully adjusted models are summarized in table 6.16. Overall the LGM explained between 17.8% and 36.9% out of the total variability of the initial growth status, between 15.5% and 38.9% out of the total variability of the linear slope terms (or total change), and between 21.5% and 69.2% out of the total variability of the quadratic terms. In case of weight, length, stunting and underweight a relatively higher proportion of variability was explained by the models for initial values than for the slope (i.e. overall change that occurred between two and eighteen months). There was an increasing trend in the proportion of variability of measured growth outcomes explained by the models as infants get older with minimal effect from how SRQ score was treated in the model (figure 6.1).

Table 6.15b: Proportion of total variation in the LGM parameters and measured growth outcomes explained by the model

Measured outcomes	*Model parameters*			***Months of growth measurement***					
	I	s	Q	2	4	6	9	12	18
	SRQ-20 was treated as a continuous exposure variable								
Length	0.260	0.214	-	0.210	0.206	0.321	0.347	0.460	0.701
Weight	0.324	0.164	-	0.498	0.494	0.458	0.391	0.512	0.571
length-for-age	0.238	0.204	0.215	0.324	0.489	0.337	0.489	0.574	0.737
Weight-for-age	0.299	0.380	0.356	0.581	0.552	0.454	0.417	0.519	0.800
Stunting	0.279	0.279		0.447	0.399	0.271	0.522	0.735	0.505
Underweight	0.178	0.342		0.489	0.406	0.569	0.496	0.995	0.628
	SRQ-20 was treated as a binary exposure variable								
Length	0.252	0.211	-	0.209	0.206	0.321	0.348	0.461	0.701
Weight	0.321	0.155	-	0.499	0.496	0.458	0.391	0.511	0.571
length-for-age	0.294	0.289	0.692	0.246	0.458	0.310	0.423	0.514	0.616
Weight-for-age	0.295	0.362	0.333	0.575	0.553	0.455	0.418	0.518	0.804
Stunting	0.281	0.228		0.497	0.701	0.287	0.763	0.611	0.435
Underweight	0.369	0.169		0.493	0.427	0.584	0.487	0.994	0.629

I = Intercept term, S = Linear slope term, Q = Quadratic term

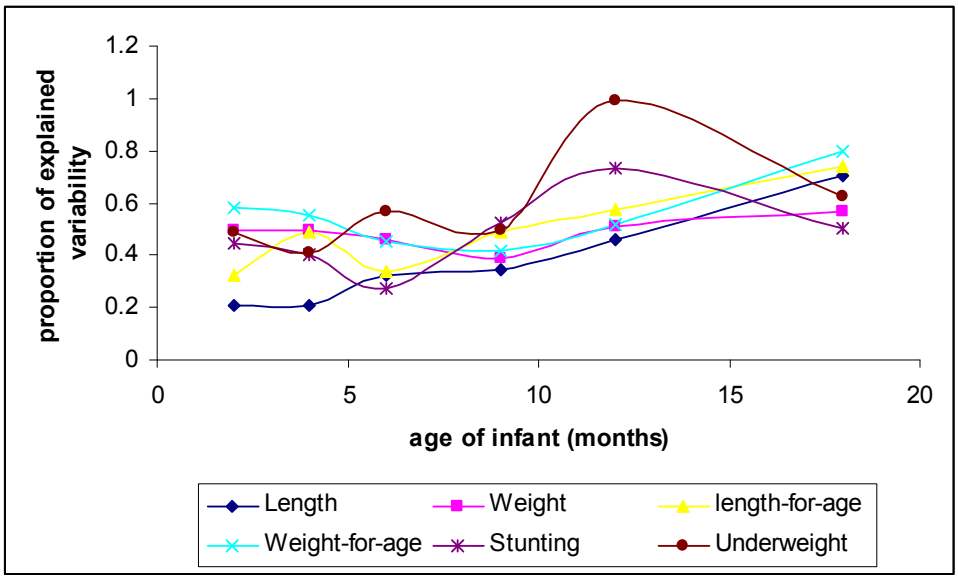

Figure 6.1: Performance of fully adjusted LGMs in explaining the variability within infant growth measurements at different follow-up time points

6.4.2 The effect of maternal CMD and other factors on infant weight

Direct effects of CMD and other risk factors on the LGM parameters of infant weight are summarized in table 6.16. Antenatal CMD was not significantly associated with initial weight of an average infant. However, it was significantly associated with small overall weight gain of an average infant between two and eighteen months of age when SRQ-20 was treated as continuous variable. When SRQ-20 was treated as a binary variable it had a negative effect on the overall weight gain of an average infant during the follow-up period but the effect size was no large enough to be statistically significant. Postnatal CMD was not significantly associated either with initial weight or the overall weight gain of an average infant during the follow-up period

Other factors associated with increased initial weight were increased maternal height, increased maternal mid-upper arm circumference, higher birth weight, having at least one obstetric complication during the delivery of index infant, and being male infant. The only factors that were significantly and inversely associated with the overall weight gain from two to eighteen months of age were increased maternal age and scoring higher on

poor sanitary condition scale of the household. All of these associations were true whether SRQ-20 was treated as a score or as a binary variable.

Figure 6.2a and 6.2b shows possible mechanisms through which maternal CMD might potentially influence infant weight. There was significant continuity of maternal CMD from pregnancy to postnatal period. Moreover, persistent CMD was positively associated with early infant illnesses (diarrhoea and life threatening illness). However, the negative effect of persistent CMD on infant illness was not translated into infant weight.

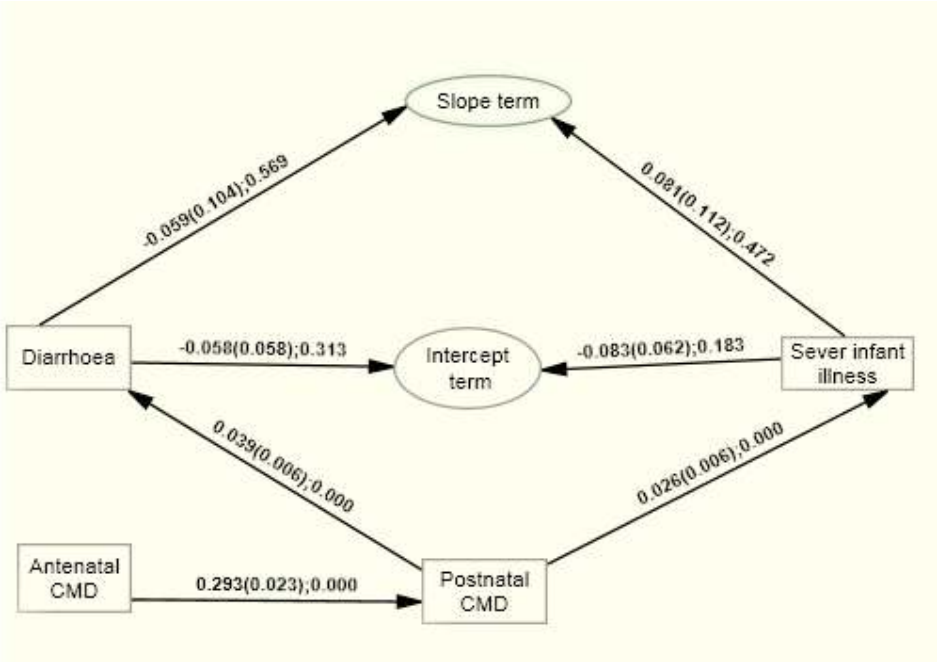

Figure 6.2a. The effect of persistent CMD on weight (SRQ-20 as continuous variable). (Numbers on the lines are *estimate(standed error); p-value*)

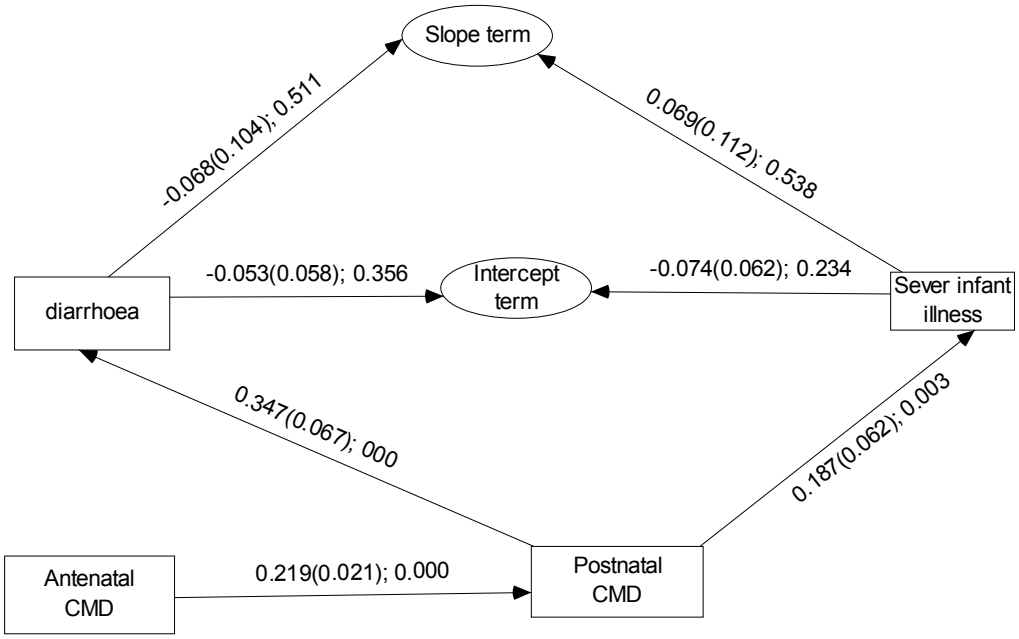

Figure 6.2b. The effect of persistent CMD on weight (SRQ-20 as binary a variable). (Numbers on the lines are *estimate(standed error); p-value*)

Table 6.16: Coefficients Estimates, standard errors and p-values for the predictors of infant weight change in the conditional LGM in Butajira, Ethiopia [**SRQ score**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ score)	0.013	0.009	0.126	-0.034	0.015	0.024
Postnatal CMD (SRQ score)	0.014	0.012	0.234	0.002	0.021	0.913
Age (years)	-0.003	0.004	0.443	-0.027	0.008	0.000
Height (cm)	0.011	0.004	0.007	-0.006	0.007	0.435
Mid upper arm circumference (cm)	0.034	0.012	0.003	0.029	0.021	0.153
Being in polygamous marriage	0.117	0.064	0.065	-0.097	0.115	0.400
Autonomy scale (0-5)						
Use khat and/or alcohol						
Had at least one obstetric complication	0.107	0.052	0.039	0.130	0.095	0.172
Household characteristics						
Urban residence	0.127	101	0.207	0.253	0.151	0.093
Number of under 5 children	-0.016	0.037	0.671	-0.042	0.063	0.524
Age of father in years						
Poverty index(0-11)						
Poor sanitary condition(0-3)	-0.030	0.030	0.321	-0.158	0.054	0.003
Level of social support(0-4)						
Characteristics of index child						
Female gender	-0.226	0.049	0.000	-0.148	0.087	0.091
Not immunized at two months						
Severe illness in the first 2 months	-0.083	0.062	0.183	0.081	0.112	0.470
Diarrhoeal episode in the first 2 months	-0.058	0.058	0.313	-0.059	0.104	0.569
Birth weight	0.592	0.075	0.000	-0.057	0.139	0.685
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	-0.030	0.068	0.662	0.085	0.122	0.485
Received pre-lacteal food	-0.048	0.121	0.691	-0.131	0.213	0.539
Colostrums given	0.017	0.061	0.784	0.198	0.112	0.076
Breast feeding delayed for 1 hours	-0.131	0.097	0.177	0.147	0.173	0.396

SE: Standard error

Table 6.17: Estimates of regression coefficients, standard errors and p-values for the predictors of infant weight change in the conditional LGM in Butajira, Ethiopia [**SRQ binary**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ binary)	0.075	0.081	0.354	-0.218	0.144	0.128
Postnatal CMD (SRQ binary)	0.182	0.124	0.142	-0.031	0.223	0.889
Age (years)	-0.003	0.004	0.536	-0.028	0.007	0.000
Height (cm)	0.011	0.004	0.005	-0.006	0.007	0.395
Mid upper arm circumference (cm)	0.035	0.012	0.003	0.029	0.021	0.160
Being in polygamous marriage	0.116	0.064	0.068	-0.099	0.115	0.388
Autonomy scale (0-5)						
Use khat and/or alcohol						
Had at least one obstetric complication	0.112	0.052	0.030	0.124	0.095	0.190
Household characteristics						
Urban residence	0.130	0.101	0.197	0.237	0.150	0.115
Number of under 5 children	-0.018	0.037	0.622	-0.039	0.066	0.553
Age of father in years						
Poverty index(0-11)						
Poor sanitary condition(0-3)	-0.028	0.030	0.362	-0.162	0.054	0.003
Level of social support(0-4)						
Characteristics of index child						
Female gender	-0.222	0.049	0.000	-0.151	0.088	0.086
Not immunized at two months						
Severe illness in the first 2 months	-0.074	0.062	0.234	0.069	0.112	0.538
Diarrhoeal episode in the first 2 months	-0.053	0.058	0.356	-0.068	0.104	0.511
Birth weight	0.594	0.075	0.000	-0.060	0.139	0.669
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	-0.030	0.068	0.662	0.081	0.122	0.507
Received pre-lacteal food	-0.055	0.121	0.651	-0.106	0.214	0.622
Colostrums given	0.014	0.062	0.820	0.199	0.112	0.076
Breast feeding delayed for 1 hours	-0.122	0.097	0.208	0.132	0.173	0.447

SE: Standard error

6.4.3 Effects of CMD and other factors on infant length

Direct effects of perinatal CMD and other risk factors on the LGM parameters of infant length are summarized in tables 6.18 and 6.19. In a fully adjusted model perinatal CMD was not significantly associated with the LGM parameters of infant length. Urban residence, higher score on poor sanitary condition of the household, being a male infant, and increased birth weight were significantly associated with an increased initial length. However, increased maternal autonomy as reported by the mother, and providing colostrums to an index infant were significantly associated with shorter initial length. Increased maternal age, scoring higher on poor sanitary condition scale of the household, scoring lower on maternal autonomy scale and not having at least one obstetric complication were significantly associated with shorter overall gain in length between two and eighteen months of age. All these associations were true regardless of how SRQ-20 score was treated within the model (as a dimensional measure or as a binary variable).

Figure 6.3a and 6.3b shows possible mechanisms through which maternal CMD might potential influence infant length. Regardless of significant continuity of maternal CMD from pregnancy to postnatal period and significant positive effect of persistent CMD on early infant illnesses (diarrhoea and life threatening illness), there was no significant association between persistent CMD and infant length.

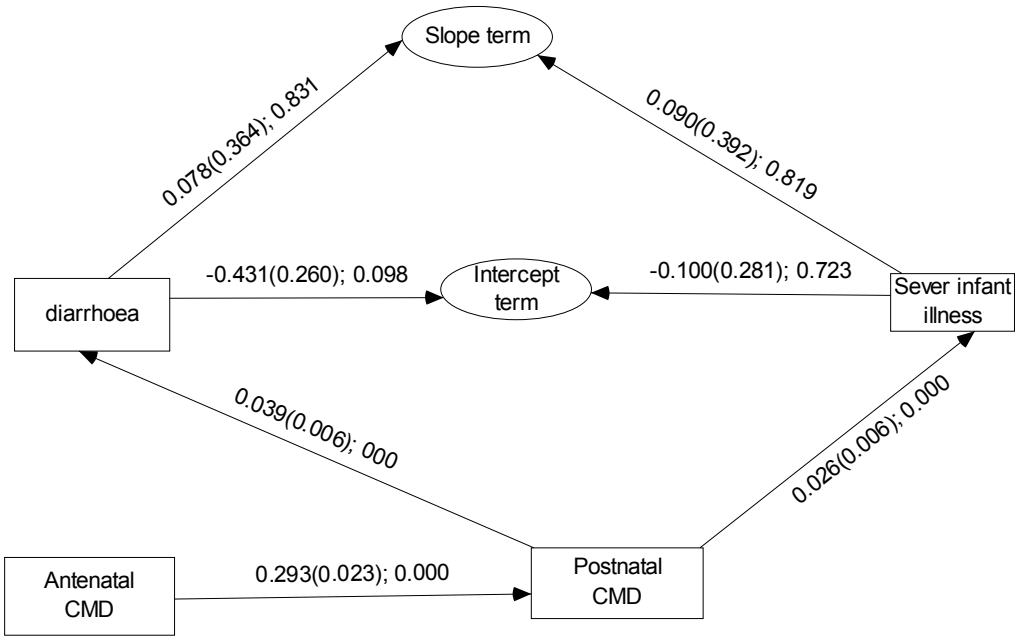

Figure 6.3a: The effect of persistent CMD on length (SRQ-20 as continuous variable). (Numbers on the lines are *estimate(standed error); p-value*)

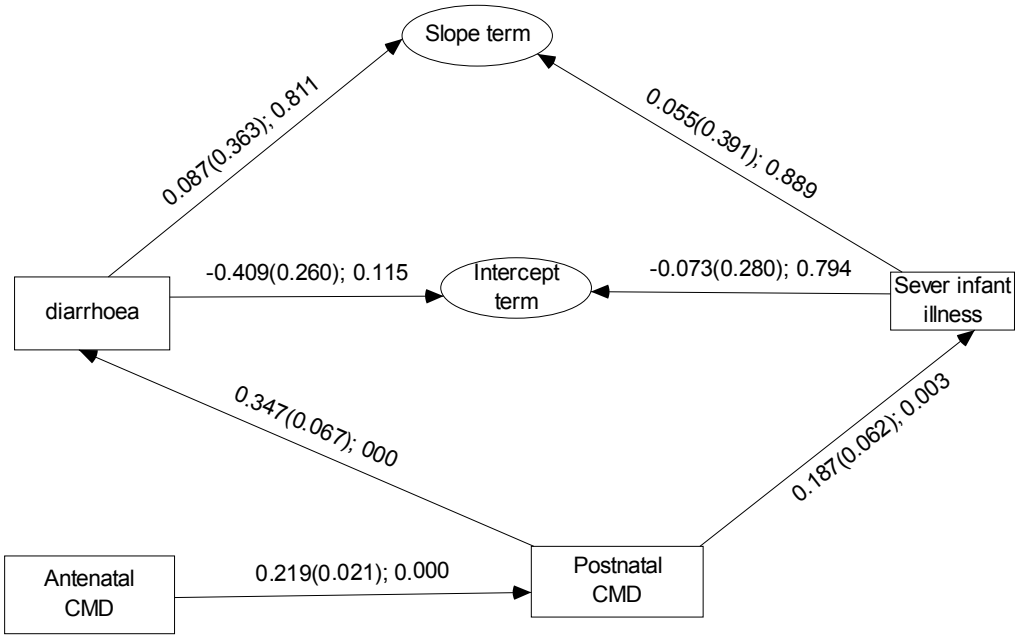

Figure 6.3b: The effect of persistent CMD on length (SRQ-20 as a binary variable). (Numbers on the lines are *estimate(standed error); p-value*)

Table 6.18: Coefficients, standard errors and p-values for the predictors of infant length change in the conditional LGM in Butajira, Ethiopia [**SRQ score**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ score)	0.065	0.038	0.087	-0.068	0.053	0.197
Postnatal CMD (SRQ score)	-0.025	0.053	0.631	0.002	0.073	0.979
Age (years)	0.028	0.018	0.133	-0.083	0.026	0.001
Height (cm)	0.005	0.018	0.783	0.021	0.025	0.394
Mid upper arm circumference (cm)	0.066	0.052	0.202	0.006	0.072	0.934
Being in polygamous marriage						
Autonomy scale (0-5)	-0.193	0.060	0.001	0.224	0.084	0.008
Use khat and/or alcohol						
Had at least one obstetric complication	-0.370	0.236	0.116	0.679	0.331	0.040
Household characteristics						
Urban residence	1.140	0.392	0.004	0.376	0.534	0.481
Number of under 5 children	-0.284	0.163	0.082	0.151	0.229	0.520
Age of father in years						
Poverty index(0-11)						
Poor sanitary condition(0-3)	0.371	0.135	0.006	-0.854	0.189	0.000
Level of social support(0-4)						
Characteristics of index child						
Female gender	-0.650	0.219	0.003	-0.126	0.306	0.680
Not immunized at two months						
Severe illness in the first 2 months	-0.100	0.281	0.723	0.090	0.392	0.819
Diarrhoeal episode in the first 2 months	-0.431	0.260	0.098	0.078	0.364	0.831
Birth weight	1.088	0.357	0.002	0.425	0.516	0.409
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months						
Received pre-lacteal food	-0.423	0.541	0.434	-1.066	0.746	0.153
Colostrums given	-0.591	0.282	0.036	0.155	0.394	0.694
Breast feeding delayed for 1 hours	0.288	0.446	0.518	0.231	0.616	0.708

SE: Standard error

Table 6.19: Coefficients, standard errors and p-values for the predictors of infant length change in the conditional LGM in Butajira, Ethiopia [**SRQ binary**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ binary)	0.393	0.361	0.270	-0.201	0.503	0.690
Postnatal CMD (SRQ binary)	-0.273	0.560	0.626	-0.491	0.781	0.530
Age (years)	0.029	0.018	0.108	-0.086	0.026	0.000
Height (cm)	0.006	0.018	0.751	0.020	0.025	0.434
Mid upper arm circumference (cm)	0.065	0.052	0.207	0.003	0.072	0.963
Being in polygamous marriage						
Autonomy scale (0-5)	-0.192	0.060	0.001	0.221	0.084	0.009
Use khat and/or alcohol						
Had at least one obstetric complication	-0.355	0.235	0.131	0.666	0.330	0.044
Household characteristics						
Urban residence	1.177	0.392	0.003	0.333	0.533	0.532
Number of under 5 children	-0.289	0.164	0.078	0.161	0.229	0.483
Age of father in years						
Poverty index(0-11)						
Poor sanitary condition(0-3)	0.380	0.135	0.005	-0.877	0.189	0.000
Level of social support(0-4)						
Characteristics of index child						
Female gender	-0.647	0.220	0.003	-0.147	0.307	0.631
Not immunized at two months						
Severe illness in the first 2 months	-0.073	0.280	-0.794	0.054	0.391	0.889
Diarrhoeal episode in the first 2 months	-0.409	0.260	0.115	0.087	0.363	0.811
Birth weight	1.117	0.357	0.002	0.386	0.516	0.455
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months						
Received pre-lacteal food	-0.455	0.544	0.403	-1.058	0.750	0.158
Colostrums given	-0.589	0.282	0.037	0.156	0.395	0.693
Breast feeding delayed for 1 hours	0.319	0.445	0.473	0.179	0.616	0.771

SE: Standard error

6.4.4 Effects of maternal CMD and other factors on infant weight-for-age z

Direct effects of perinatal CMD and other risk factors on the LGM parameters of infant weight-for-age are summarized in tables 6.20 and 16.21. Postnatal CMD was not significantly associated with the LGM parameters of weight-for-age z. When SRQ-20 was treated as a continuous variable antenatal CMD was significantly associated with higher initial value and inversely associated with the linear term of weight-for-age z.

Regardless of how SRQ-20 score was treated in the LGM an increase in the height of the mother, being a female infant and an increase in birth weight were significantly associated with increased initial weight-for-age z of an average infant. Higher score on poor sanitary condition scale, an increased birth weight, and an increased maternal age were significantly and inversely associated with linear slope term and positively associated with quadratic term of weight-for-age z score whether SRQ-20 was treated as a continuous or as a binary variable. None of the early infant feeding practices of the mother were significantly associated with the LGM parameters of weight-for-age z Significant continuity of maternal CMD from pregnancy to postnatal period and positive association of postnatal CMD with early infant illness was not translated to the negative effect of persistent CMD on a reduced weight-for-age z of an average infant mediated through infant illnesses (figure 6.4a and figure 6.4b). This negative association between persistent CMD and weight-for-age z was independent of the scale used for SRQ-20 within the LGM of weight-for-age.

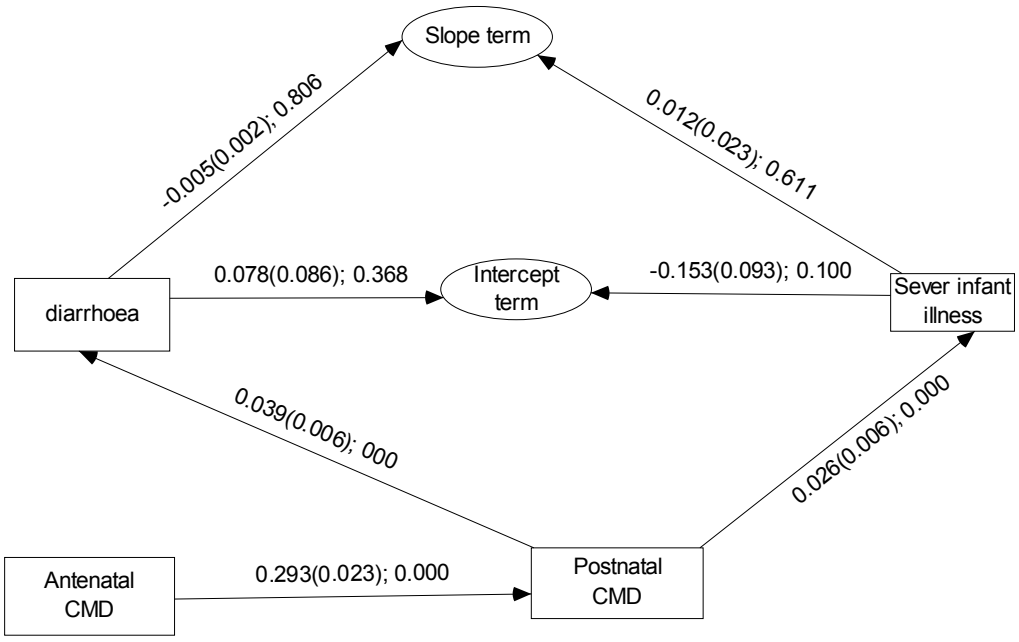

Figure 6.4a: The effect of persistent CMD on weight-for-age z (SRQ-20 as continuous variable). (Numbers on the lines are *estimate(standed error); p-value*)

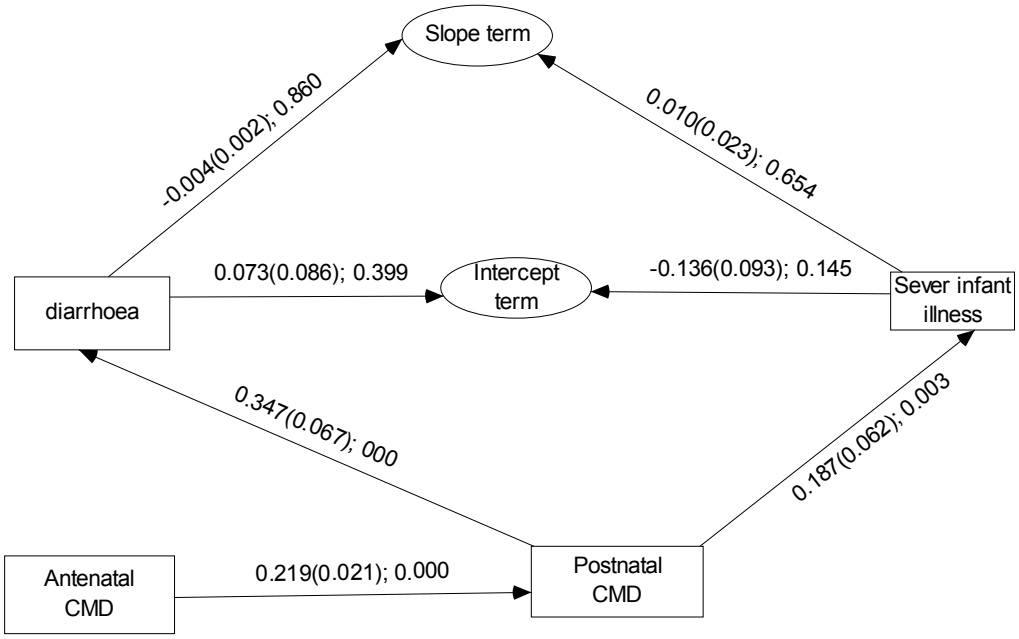

Figure 6.4b: The effect of persistent CMD on weight-for-age z (SRQ-20 as binary variable). (Numbers on the lines are *estimate(standed error); p-value*)

Table 6.20: Coefficients Estimates, standard errors and p-values for the predictors of infant weight-for-age change in the conditional LGM in Butajira, Ethiopia [**SRQ as a score**]

Predictor variable	Intercept factor			***Slope factor***			*Quadratic factor*		
	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother									
Antenatal CMD (SRQ score)	0.029	0.013	0.021	-0.008	0.003	0.010	0.000	0.000	0.052
Postnatal CMD (SRQ score)	0.012	0.017	0.481	0.006	0.004	0.164	0.000	0.000	0.074
Age (years)	-0.003	0.006	0.621	-0.004	0.002	0.011	0.000	0.000	0.048
Height (cm)	0.016	0.006	0.007	-0.002	0.002	0.203	0.000	0.000	0.354
Mid upper arm circumference (cm)	0.033	0.017	0.057	0.008	0.004	0.055	0.000	0.000	0.048
Being in polygamous marriage									
Autonomy scale (0-5)									
Use khat and/or alcohol									
Had at least one obstetric complication	0.141	0.078	0.069	0.019	0.020	0.332	-0.001	0.001	0.351
Household characteristics									
Urban residence	0.220	0.158	0.164	0.049	0.033	0.138	-0.003	0.002	0.142
Number of under 5 children	0.002	0.055	0.964	-0.015	0.014	0.275	0.001	0.001	0.298
Age of father in years									
Poverty index(0-11)									
Poor sanitary condition(0-3)	-0.021	0.046	0.651	-0.031	0.011	0.007	0.001	0.001	0.025
Level of social support(0-4)									
Characteristics of index child									
Female gender	0.342	0.073	0.000	0.024	0.018	0.185	-0.002	0.001	0.083
Not immunized at two months									
Severe illness in the first 2 months	-0.153	0.093	0.100	0.012	0.023	0.611	0.000	0.001	0.981
Diarrhoeal episode in the first 2 months	-0.078	0.086	0.368	-0.005	0.022	0.806	0.000	0.001	0.849
Birth weight	1.023	0.111	0.000	-0.111	0.028	0.000	0.005	0.002	0.000
Feeding practices in the first two months of infancy									
Non-exclusive breast-feeding at 2 months									
Received pre-lacteal food	-0.013	0.181	0.942	-0.057	0.045	0.204	0.003	0.003	0.177
Colostrums given	0.001	0.092	0.988	0.025	0.023	0.278	-0.001	0.001	0.495
Breast feeding delayed for 1 hours	-0.088	0.145	0.547	0.001	0.036	0.976	0.000	0.002	0.844

SE: Standard error

Table 6.21: Coefficients Estimates, standard errors and p-values for the predictors of infant weight-for-age change in the conditional LGM in Butajira, Ethiopia [**SRQ as binary exposure**]

Predictor variable	Intercept factor			Slope factor			Quadratic factor		
	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother									
Antenatal CMD (SRQ binary)	0.127	0.120	0.292	-0.045	0.030	0.130	0.002	0.002	0.258
Postnatal CMD (SRQ binary)	0.250	0.184	0.174	0.016	0.046	0.732	-0.002	0.003	0.524
Age (years)	-0.002	0.006	0.764	-0.004	0.002	0.007	0.000	0.000	0.040
Height (cm)	0.017	0.006	0.005	-0.002	0.002	0.181	0.000	0.000	0.340
Mid upper arm circumference (cm)	0.034	0.018	0.051	0.008	0.004	0.056	0.000	0.000	0.048
Being in polygamous marriage									
Autonomy scale (0-5)									
Use khat and/or alcohol									
Had at least one obstetric complication	0.149	0.078	0.056	0.020	0.020	0.305	-0.001	0.001	0.298
Household characteristics									
Urban residence	0.234	0.158	0.139	0.044	0.033	0.183	-0.002	0.002	0.179
Number of under 5 children	-0.002	0.055	0.973	-0.014	0.014	0.303	0.001	0.001	0.320
Age of father in years									
Poverty index(0-11)									
Poor sanitary condition(0-3)	-0.014	0.046	0.755	-0.032	0.011	0.005	0.001	0.001	0.020
Level of social support(0-4)									
Characteristics of index child									
Female gender	0.349	0.074	0.000	0.023	0.018	0.206	-0.002	0.001	0.092
Not immunized at two months									
Severe illness in the first 2 months	-0.136	0.093	0.145	0.010	0.023	0.654	0.000	0.001	0.977
Diarrhoeal episode in the first 2 months	-0.073	0.086	0.399	-0.004	0.022	0.860	0.000	0.001	0.954
Birth weight	1.026	0.111	0.000	-0.112	0.028	0.000	0.006	0.002	0.000
Feeding practices in the first two months of infancy									
Non-exclusive breast-feeding at 2 months									
Received pre-lacteal food	-0.026	0.182	0.884	-0.052	0.045	0.248	0.003	0.003	0.205
Colostrums given	-0.003	0.092	0.977	0.025	0.023	0.277	-0.001	0.001	0.503
Breast feeding delayed for 1 hours	-0.066	0.145	0.651	-0.001	0.036	0.986	0.000	0.002	0.845

6.4.5 Effects of maternal CMD and other factors on infant length-for-age

Direct effects of perinatal CMD and other risk factors on the LGM parameters of infant length-for-age are summarized in tables 6.22 and 16.23. Antenatal CMD had a significant positive effect on initial value and quadratic term and significant negative effect on linear slope term of the LGM underlying length-for-age z of an average infant. Similarly, postnatal CMD had significant positive effect on linear slope term and significant negative effect on quadratic term of the LGM of length-for-age z of an average infant. These associations were independent of whether SRQ-20 was treated as a binary or as a continuous variable in the LGM. The effect of postnatal CMD on initial value of length-for-age z was dependent on how SRQ-20 was treated in the model – it had significant negative effect when SRQ-20 was included in the model as a continuous variable and non-significant, but still with the same sign, when it was used as a binary exposure variable. There was a significant negative effect of persistent CMD through infant diarrhoea on linear slope, but not on the initial value, of LGM underlying length-for-age z of average infant. There was a non-significant trend of the negative effect of persistent CMD on the initial value through severe life threatening infant illness **(**figures 6.5a and 6.5b**)**.

Rural residence and withholding of colostrums had a significant negative effect on initial value of LGM underlying length-for-age z but they were not significantly associated with its linear or quadratic terms. Diarrhoeal episodes in the first two months of infancy and delayed initiation of breastfeeding were not significantly associated with the initial value of LGM underlying length-for-age z but they had a significant positive effect on the linear slope term and a significant negative effect on the quadratic term of the LGM. Higher score on poor sanitary condition scale of the household and increased value of birth weight had significant positive effect on initial value and significant negative effect on linear slope of LGM underlying length-for-age z of an average infant. Female gender had a significant positive effect on initial value and linear slope of the LGM underlying length-for-age z. All these associations were independent of whether SRQ-20 was treated as a continuous or as a binary variable within the model. Regardless of how SRQ-20 was treated in the model maternal MUAC was not significantly associated with initial value

of length-for-age but it had marginally significant positive effect on the linear slope when SRQ-20 was treated as a binary exposure variable.

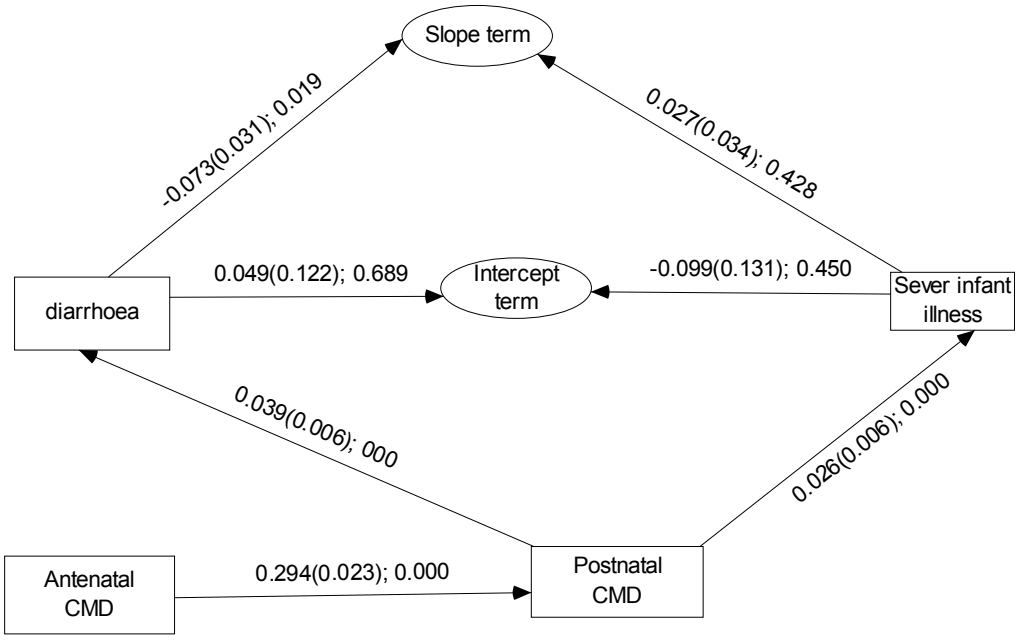

Figure 6.5a: The effect of persistent CMD on length-for-age z (SRQ-20 as continuous variable). (Numbers on the lines are *estimate(standed error); p-value*)

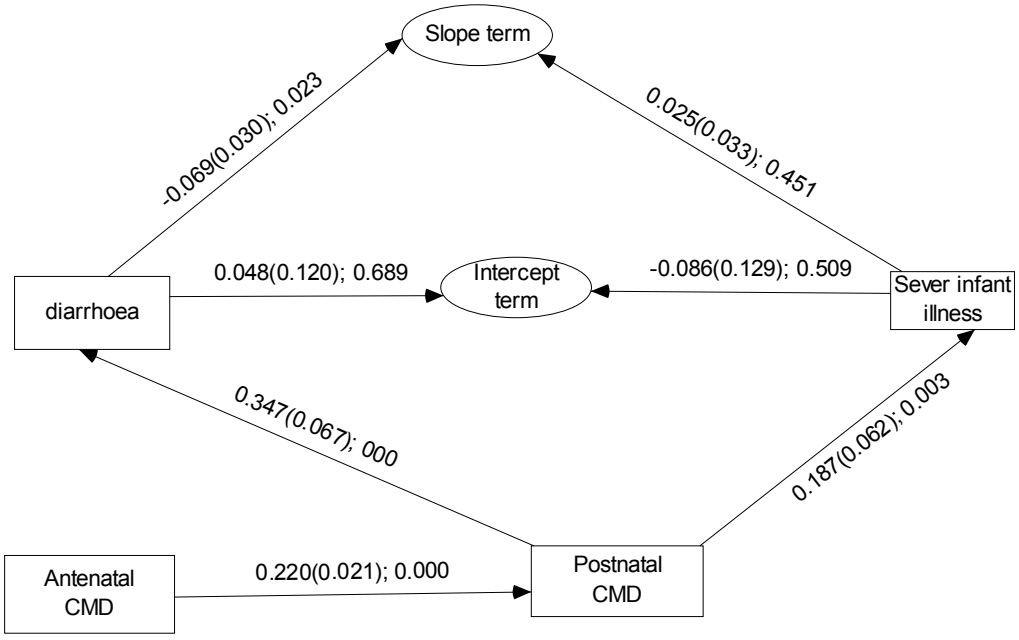

Figure 6.5b: The effect of persistent CMD on length-for-age z (SRQ-20 as a binary variable). (Numbers on the lines are *estimate(standed error); p-value*)

Table 6.22: Coefficients Estimates, standard errors and p-values for the predictors of infant length-for-age change in the conditional LGM in Butajira, Ethiopia [**SRQ as a score**]

Predictor variable	Intercept factor			Slope factor			Quadratic factor		
	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother									
Antenatal CMD (SRQ score)	0.062	0.018	0.000	-0.017	0.005	0.000	0.001	0.000	0.000
Postnatal CMD (SRQ score)	-0.061	0.025	0.014	0.022	0.006	0.000	-0.001	0.000	0.000
Age (years)	-0.012	0.012	0.329	0.000	0.003	0.938	0.000	0.000	0.553
Height (cm)	0.011	0.008	0.216	-0.002	0.002	0.429	0.000	0.000	0.283
Mid upper arm circumference (cm)	0.002	0.024	0.928	0.012	0.006	0.062	-0.001	0.000	0.040
Being in polygamous marriage									
Autonomy scale (0-5)									
Use khat and/or alcohol									
Had at least one obstetric complication	-0.109	0.110	0.321	0.030	0.028	0.291	-0.001	0.001	0.409
Household characteristics									
Urban residence	0.632	0.191	0.000	-0.021	0.046	0.654	0.001	0.002	0.561
Number of under 5 children	-0.057	0.078	0.460	-0.012	0.019	0.534	0.001	0.001	0.449
Age of father in years	0.009	0.008	0.285	0.001	0.002	0.706	0.000	0.000	0.564
Poverty index(0-11)									
Poor sanitary condition(0-3)	0.153	0.064	0.016	-0.046	0.016	0.004	0.002	0.001	0.051
Level of social support(0-4)									
Characteristics of index child									
Female gender	0.351	0.103	0.000	0.053	0.026	0.044	-0.003	0.001	0.031
Not immunized at two months									
Severe illness in the first 2 months	-0.099	0.131	0.450	0.027	0.034	0.428	-0.001	0.002	0.479
Diarrhoeal episode in the first 2 months	0.049	0.122	0.689	-0.073	0.031	0.019	0.004	0.002	0.011
Birth weight	0.722	0.169	0.000	-0.089	0.044	0.042	0.005	0.002	0.021
Feeding practices in the first two months of infancy									
Non-exclusive breast-feeding at 2 months	-0.181	0.140	0.196	0.029	0.036	0.425	-0.001	0.002	0.578
Received pre-lacteal food	-0.347	0.255	0.173	0.001	0.064	0.982	-0.001	0.003	0.731
Colostrums given	-0.301	0.131	0.022	0.025	0.034	0.459	-0.001	0.002	0.494
Breast feeding delayed for 1 hours	-0.224	0.209	0.284	0.125	0.053	0.018	-0.007	0.003	0.017

Table 6.23: Coefficients Estimates, standard errors and p-values for the predictors of infant length-for-age change in the conditional LGM in Butajira, Ethiopia [**SRQ as binary exposure**]

Predictor variable	Intercept factor			Slope factor			Quadratic factor		
	Estimate	SE	p-value	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother									
Antenatal CMD (SRQ binary)	0.382	0.166	0.022	-0.102	0.042	0.014	0.005	0.002	0.014
Postnatal CMD (SRQ binary)	-0.407	0.259	0.117	0.136	0.065	0.037	-0.009	0.003	0.012
Age (years)	-0.012	0.012	0.308	0.001	0.003	0.829	0.000	0.000	0.526
Height (cm)	0.012	0.008	0.168	-0.002	0.002	0.353	0.000	0.000	0.281
Mid upper arm circumference (cm)	0.001	0.024	0.962	0.012	0.006	0.044	-0.001	0.000	0.024
Being in polygamous marriage									
Autonomy scale (0-5)									
Use khat and/or alcohol									
Had at least one obstetric complication	-0.119	0.108	0.271	0.035	0.027	0.198	-0.002	0.001	0.284
Household characteristics									
Urban residence	0.682	0.190	0.000	-0.040	0.045	0.379	0.002	0.002	0.297
Number of under 5 children	-0.067	0.077	0.387	-0.010	0.019	0.599	0.001	0.001	0.500
Age of father in years	0.010	0.008	0.224	0.000	0.002	0.871	0.000	0.000	0.721
Poverty index(0-11)									
Poor sanitary condition(0-3)	0.162	0.063	0.010	-0.050	0.016	0.002	0.002	0.001	0.025
Level of social support(0-4)									
Characteristics of index child									
Female gender	0.362	0.102	0.000	0.051	0.026	0.046	-0.003	0.001	0.030
Not immunized at two months									
Severe illness in the first 2 months	-0.086	0.129	0.509	0.025	0.033	0.451	-0.001	0.002	0.487
Diarrhoeal episode in the first 2 months	0.048	0.120	0.689	-0.069	0.030	0.023	0.004	0.002	0.014
Birth weight	0.752	0.167	0.000	-0.096	0.043	0.025	0.006	0.002	0.012
Feeding practices in the first two months of infancy									
Non-exclusive breast-feeding at 2 months	-0.169	0.138	0.223	0.028	0.035	0.435	-0.001	0.002	0.597
Received pre-lacteal food	-0.385	0.254	0.130	0.013	0.063	0.838	-0.002	0.003	0.591
Colostrums given	-0.303	0.130	0.020	0.026	0.033	0.418	-0.001	0.002	0.449
Breast feeding delayed for 1 hours	-0.230	0.206	0.264	0.128	0.051	0.012	-0.007	0.003	0.011

6.4.6 Effects of maternal CMD and other factors on the probability of stunting

Direct effects of perinatal CMD on the LGM parameters of stunting (i.e. on a probit scale) are summarized in Tables 6.24 and 6.25. Independent of how SRQ-20 was treated in the model antenatal CMD was significantly associated with increased risk of stunting at baseline but it was not significantly associated with the overall change of the probit of stunting during the follow-up time. When SRQ-20 was treated as a dimensional measure postnatal CMD was significantly associated with reduced risk of baseline stunting and inversely associated with the overall change in the probit of stunting over the follow-up time. Antenatal CMD had no significant effect on infant stunting mediated through early infant feeding practices of the mother or through birth weight.

Direct and indirect effects of persistent CMD on stunting mediated through early infant illness are summarized in Table 6.26. Persistent CMD was not significantly associated with stunting when SRQ-20 was treated as a binary variable. However, it had a non-mediated significant positive association with the probit of baseline stunting ($\hat{\beta}$ = 0.016, SE = 0.008, p-value = 0.044) and a significant negative association with the probit of overall change of stunting between two and eighteen months ($\hat{\beta}$ = -0.016, SE = 0.008, p-value = 0.030) when SRQ-20 was treated as a continuous variable. Moreover, it had a marginally non-significant positive association with the probit of overall change of stunting between two and eighteen months of age which was mediated through infant diarrhoea ($\hat{\beta}$ = 0.004, SE = 0.002, p-value = 0.059). There was no significant effect of persistent CMD on the probits of infant stunting that was mediated though life threatening sever illness of infants of any sort in the first two months of life.

The effects of selected risk factors on the LGM parameters of stunting (i.e. on a probit scale) are summarized in Tables 6.24 and 6.25. Urban residence, female gender of the infant and an increase in birth weight were significantly and inversely associated with initial probit of stunting (i.e. they were predictors of reduced baseline probability of stunting). Similarly, delayed initiation of breast feeding for more than an hours and scoring lower on poor sanitary condition scale of the household were inversely associated with the overall change in the probit of stunting between two and eighteen months of age. All these associations were independent from how SRQ-20 was treated in the model.

History of diarrhoeal episodes in the first two months of age was positively associated with the overall change in the probit of stunting between two and eighteen months of age only when SRQ-20 was treated as continuous variable and it was not associated with the initial level of the probit regardless of how SRQ-20 was treated in the model.

Table 6.24: Coefficients Estimates, standard errors and p-values for the predictors of infant stunting in the conditional LGM in Butajira, Ethiopia [**SRQ as a binary**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ - binary)	-0.703	0.320	0.028	0.447	0.269	0.096
Postnatal CMD (SRQ - binary)	0.260	0.153	0.089	-0.182	0.127	0.152
Age (years)	0.007	0.016	0.677	0.004	0.011	0.722
Height (cm)	-0.014	0.011	0.232	0.006	0.008	0.455
Mid upper arm circumference (cm)	-0.011	0.033	0.736	-0.021	0.023	0.355
Being in polygamous marriage	-0.099	0.196	0.614	0.123	0.134	0.358
Autonomy scale (0-5)	0.030	0.038	0.422	-0.011	0.026	0.687
Use khat and/or alcohol	---		---		--	
Had at least one obstetric complication	0.000	0.139	0.999	-0.021	0.095	0.825
Household characteristics						
Urban residence	-1.041	0.394	0.008	0.272	0.314	0.387
Number of under 5 children	0.088	0.104	0.397	0.001	0.072	0.989
Age of father in years	0.008	0.013	0.529	-0.009	0.009	0.334
Poverty index(0-11)	-0.050	0.048	0.299	0.015	0.036	0.675
Poor sanitary condition(0-3)	-0.107	0.086	0.210	0.150	0.072	0.036
Level of social support(0-4)	0.049	0.062	0.435	-0.030	0.044	0.502
Characteristics of index child						
Female gender	-0.500	0.152	0.001	-0.075	0.113	0.509
Not immunized at two months						
Severe illness in the first 2 months	-0.069	0.096	0.474	-0.012	0.070	0.861
Diarrhoeal episode in the first 2 months	-0.043	0.109	0.691	0.150	0.081	0.062
Birth weight	-0.703	0.220	0.001	0.066	0.170	0.700
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	-0.118	0.187	0.530	0.138	0.127	0.280
Received pre-lacteal food	0.210	0.247	0.397	0.064	0.173	0.709
Colostrums given	0.151	0.172	0.381	-0.072	0.114	0.527
Breast feeding delayed for 1 hours	0.084	0.093	0.369	-0.166	0.074	0.025

Table 6.25: Coefficients Estimates, standard errors and p-values for the predictors of infant stunting in the conditional LGM in Butajira, Ethiopia [**SRQ as a score**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ as a score)	-0.065	0.028	0.019	0.046	0.027	0.092
Postnatal CMD (SRQ as a score)	0.055	0.027	0.043	-0.055	0.025	0.028
Age (years)	0.008	0.016	0.597	0.003	0.010	0.792
Height (cm)	-0.013	0.011	0.237	0.002	0.008	0.777
Mid upper arm circumference (cm)	-0.014	0.033	0.664	-0.022	0.021	0.297
Being in polygamous marriage	-0.142	0.193	0.463	0.142	0.129	0.270
Autonomy scale (0-5)	0.026	0.037	0.472	-0.007	0.025	0.785
Use khat and/or alcohol						
Had at least one obstetric complication	0.076	0.136	0.577	-0.057	0.089	0.520
Household characteristics						
Urban residence	-1.050	0.426	0.014	0.101	0.419	0.809
Number of under 5 children	0.044	0.097	0.650	0.038	0.064	0.557
Age of father in years	0.006	0.012	0.622	-0.007	0.008	0.414
Poverty index(0-11)	-0.042	0.047	0.377	-0.002	0.035	0.960
Poor sanitary condition(0-3)	-0.137	0.089	0.124	0.162	0.077	0.036
Level of social support(0-4)	0.046	0.060	0.449	-0.015	0.040	0.709
Characteristics of index child						
Female gender	-0.412	0.160	0.010	-0.235	0.163	0.150
Not immunized at two months						
Severe illness in the first 2 months	-0.052	0.099	0.597	-0.039	0.069	0.572
Diarrhoeal episode in the first 2 months	0.019	0.098	0.849	0.131	0.067	0.052
Birth weight	-0.511	0.207	0.013	-0.204	0.194	0.294
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	-0.180	0.193	0.351	0.172	0.130	0.188
Received pre-lacteal food	0.133	0.129	0.303	0.040	0.096	0.675
Colostrums given	0.084	0.157	0.594	-0.021	0.100	0.833
Breast feeding delayed for 1 hours	0.067	0.101	0.506	-0.179	0.076	0.019

Table 6.26: Fully adjusted direct and indirect effects of persistent CMD on infant stunting

Scale of SRQ (binary or continuous)	*Latent Growth Model parameter*	*Mediator*	*Direct and indirect effects of persistent CMD*		
			Estimate	SE	p-value
Binary					
	Intercept or initial value	None	0.305	0.183	0.095
		Diarrhoea	-0.018	0.046	0.698
		Sever infant illness	-0.017	0.027	0.511
	Slope or overall change between 2 and 18 months of age	None	-0.214	0.151	0.156
		Diarrhoea	0.062	0.039	0.115
		Sever infant illness	-0.003	0.018	0.862
Continuous					
	Intercept or initial value	None	0.016	0.008	0.044
		Diarrhoea	0.001	0.003	0.849
		Sever infant illness	-0.001	0.002	0.600
	Slope or overall change between 2 and 18 months of age	None	-0.016	0.008	0.030
		Diarrhoea	0.004	0.002	0.059
		Sever infant illness	-0.001	0.002	0.575

6.4.7 Effects of maternal CMD and other factors on the probability of infant underweight

The effects of perinatal CMD on LGM parameters of infant underweight (i.e. probit scores) are summarized in Tables 6.27 and 6.28. Maternal CMD was not significantly associated with the probability of underweight independent of whether SRQ-20 was include in the models as a binary or as a continuous predictor variable.

Antenatal CMD had no significant effect on infant underweight mediated through early infant feeding practices of the mother or through birth weight. Direct and indirect effects of persistent CMD on underweight mediated through early infant illness are summarized in Table 6.29. Persistent CMD had no direct effect or indirect effects through infant illness in the first two months of life on underweight independent of whether SRQ-20 was used as a binary variable or as a continuous variable.

The effects of selected risk factors on LGM parameters of infant underweight (i.e. probit scores) are summarized in Tables 6.27 and 6.28. Factors associated with reduced risk of baseline level of infant underweight were scoring lower on poor sanitary condition scale of the household, good maternal nutritional status (i.e. increased maternal mid-upper arm circumference), being a female infant, increased birth weight and having a mother who had at least one obstetric complication during delivery. None of the pre-specified risk factors for infant underweight had a significant effect on the overall change of the probit of underweight of infants from two to eighteen months of age. All these associations were independent from how SRQ-20 was treated in the model.

Table 6.27: Coefficients Estimates, standard errors and p-values for the predictors of infant underweight in the conditional LGM in Butajira, Ethiopia [**SRQ as a score**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ as a score)	-0.018	0.021	0.396	0.022	0.016	0.182
Postnatal CMD (SRQ as a score)	-0.027	0.023	0.230	-0.015	0.015	0.312
Age (years)	0.017	0.014	0.229	0.009	0.009	0.340
Height (cm)	-0.015	0.009	0.096	0.002	0.008	0.826
Mid upper arm circumference (cm)	-0.084	0.031	0.007	-0.025	0.027	0.356
Being in polygamous marriage	0.099	0.170	0.559	-0.088	0.114	0.438
Autonomy scale (0-5)	-0.050	0.036	0.163	0.022	0.026	0.388
Use khat and/or alcohol						
Had at least one obstetric complication	-0.255	0.128	0.047	0.050	0.104	0.630
Household characteristics						
Urban residence	0.015	0.260	0.954	-0.270	0.169	0.109
Number of under 5 children	-0.106	0.092	0.252	0.093	0.066	0.155
Age of father in years	-0.004	0.011	0.707	0.006	0.006	0.326
Poverty index(0-11)	0.037	0.047	0.432	-0.047	0.035	0.172
Poor sanitary condition(0-3)	0.200	0.073	0.006	0.093	0.062	0.132
Level of social support(0-4)	0.005	0.053	0.926	0.025	0.034	0.463
Characteristics of index child						
Female gender	-0.613	0.170	0.000	0.048	0.185	0.794
Not immunized at two months						
Severe illness in the first 2 months	0.161	0.132	0.221	0.053	0.091	0.563
Diarrhoeal episode in the first 2 months	0.163	0.120	0.172	0.026	0.081	0.750
Birth weight	-0.753	0.176	0.000	-0.177	0.207	0.394
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	0.078	0.159	0.624	-0.080	0.103	0.434
Received pre-lacteal food	0.043	0.115	0.710	0.021	0.079	0.793
Colostrums given	-0.031	0.082	0.702	-0.046	0.053	0.386
Breast feeding delayed for 1 hours	0.066	0.095	0.486	-0.094	0.071	0.185

Table 6.28: Coefficients Estimates, standard errors and p-values for the predictors of infant underweight in the conditional LGM in Butajira, Ethiopia [**SRQ as binary**]

Predictor variable	Intercept factor			Slope factor		
	Estimate	SE	p-value	Estimate	SE	p-value
Characteristics of mother						
Antenatal CMD (SRQ as binary)	0.179	0.250	0.472	0.011	0.181	0.953
Postnatal CMD (SRQ as binary)	-0.201	0.147	0.171	-0.024	0.110	0.825
Age (years)	0.015	0.014	0.291	0.011	0.009	0.249
Height (cm)	-0.016	0.009	0.091	0.002	0.008	0.768
Mid upper arm circumference (cm)	-0.084	0.032	0.008	-0.025	0.027	0.359
Being in polygamous marriage	0.098	0.171	0.566	-0.098	0.121	0.417
Autonomy scale (0-5)	-0.051	0.037	0.169	0.025	0.027	0.346
Use khat and/or alcohol						
Had at least one obstetric complication	-0.265	0.132	0.046	0.066	0.110	0.552
Household characteristics						
Urban residence	-0.013	0.263	0.961	-0.266	0.173	0.124
Number of under 5 children	-0.110	0.094	0.241	0.104	0.069	0.129
Age of father in years	-0.004	0.011	0.710	0.007	0.007	0.316
Poverty index(0-11)	0.034	0.046	0.460	-0.049	0.035	0.165
Poor sanitary condition(0-3)	0.191	0.073	0.009	0.100	0.062	0.105
Level of social support(0-4)	0.000	0.054	0.999	0.029	0.036	0.415
Characteristics of index child						
Female gender	-0.616	0.170	0.000	0.062	0.185	0.737
Not immunized at two months						
Severe illness in the first 2 months	0.178	0.141	0.206	0.077	0.103	0.457
Diarrhoeal episode in the first 2 months	0.214	0.146	0.143	0.035	0.104	0.737
Birth weight	-0.767	0.181	0.000	-0.157	0.214	0.463
Feeding practices in the first two months of infancy						
Non-exclusive breast-feeding at 2 months	0.069	0.158	0.664	-0.076	0.109	0.485
Received pre-lacteal food	0.038	0.113	0.738	0.023	0.083	0.784
Colostrums given	-0.035	0.082	0.670	-0.042	0.056	0.454
Breast feeding delayed for 1 hours	0.060	0.093	0.519	-0.089	0.071	0.215

Table 6.29: Fully adjusted direct and indirect effects of persistent CMD on infant underweight

Scale of SRQ (binary or continuous)	*Latent Growth Model parameter*	*Mediator*	*Direct and indirect effects of persistent CMD*		
			Estimate	SE	p-value
Binary					
	Intercept or initial value	None	-0.245	0.188	0.193
		Diarrhoea	0.026	0.022	0.247
		Sever infant illness	0.014	0.013	0.279
	Slope or overall change between 2 and 18 months of age	None	-0.030	0.134	0.825
		Diarrhoea	0.004	0.013	0.740
		Sever infant illness	0.006	0.008	0.480
Continuous					
	Intercept or initial value	None	-0.008	0.007	0.232
		Diarrhoea	0.002	0.002	0.222
		Sever infant illness	0.001	0.001	0.248
	Slope or overall change between 2 and 18 months of age	None	-0.004	0.004	0.313
		Diarrhoea	0.000	0.001	0.751
		Sever infant illness	0.000	0.001	0.570

CHAPTER 7: RESULTS - MULTILEVEL GROWTH MODELLING

7.1 Introduction

In this chapter we present results of multilevel growth models (MGM) for six infant growth outcome measurements (length, weight, length-for-age, weight-for-age, stunting, and underweight) that cover the first 18 months of life using data from the P-MaMiE study. The correlates of the outcome variables were selected from the three domains: maternal characteristics, household characteristics, and infant characteristics. Presentation of the findings will follow the following outline:

1. Unconditional MGM: - Identification of the best fitting level-1 growth model.

The best fitting level 1 model was identified for each growth outcome after evaluating the fit of several nested polynomial functions in a hierarchical manner

2. Conditional MGM – crude and adjusted effects of level-2 predictors of infant growth at two months of age

Having identified the best fitting unconditional growth model, unadjusted effects of level -2 predictors of growth attainment at two months of age were assessed. Regardless of the statistical significance of the crude fixed effects, all pre-specified level-2 predictors of growth attainments at the age of two months were included in a multivariable conditional MGM and their adjusted effects were investigated.

3. Conditional MGM – crude and adjusted effects of level-2 predictors of infant growth attainment at the age of two months and the rates at which these growth outcomes change over time

The usefulness of having the interaction of each level-2 predictor of initial value with the age of infant while having the same level-2 predictor as a predictor of an initial growth attainment was investigated based on the best fitting unconditional MGM. The importance of keeping a particular interaction was assessed using (a) statistical significance of the fixed part (i.e. if the regression coefficient of the interaction was statistically significant) and (b) the change in a deviance as the result of including the

interaction term. In the multivariable MGM only significant interactions (i.e. significant predictors of rate of change in the target growth outcome) were included from this step of modelling. In other words the fully adjusted models were based on the best fitting unconditional MGM and had (a) all covariates included as the level-2 predictors of growth at the age of two months irrespective of their statistical significance and (b) only covariates which turned out to be significant in unadjusted models were included as predictors of the rates of growth outcomes.

4. Conditional and unconditional MGM – The effect of CMD on initial growth outcomes and on the rate of change of growth outcomes over time

The best fitting unconditional growth models were used as the base models to obtain unadjusted effect of CMD on (a) the growth outcomes attained by infants at the age of two months and on (b) the rates of change of growth outcomes over time. The adjusted effects of CMD on each growth outcome were obtained using two separate models: (a) the fully adjusted model with all covariates as the predictors of growth only at two months of age (i.e. initial value) and (b) the fully adjusted model with all covariates as predictors of growth at two months of age and selected covariates included as predictors of rate of change of growth outcomes. The selection of covariates to be included as predictors of rate of growth for each outcome was based on result of bivariate analysis (statistical significance of the interaction terms and the level of reduction of deviance in bivariate analysis)

7.2 Unconditional MGM – Identification of the best fitting level 1 growth model

7.2.1 Gain in length of infants over the first 18 months of age

The performance of four alternative growth models in describing the trajectories of length of infants over the first 18 months of life is summarized in table 7.1. As we move up in the hierarchy of model fitting from more restrictive (model 1) to less restrictive (model 4) growth model there is always significant improvement in the model fit which can be assessed using one of the two information criteria or likelihood ratio. The mean length of infants at two months of age and within infant variability over time are reduced as we move towards less restrictive model.

Among the four models the quadratic model with random intercept and random slope fits best to the length of infants. The available data was insufficient to enable us to estimate the standard errors of the random components of quadratic component. Using results from model 4 the mean length of an average infant at the age of two months is 57.6 cm with significant variability among infants (standard deviation (sd) = 2.16cm; 95% CI: 1.94 to 2.40). At this age an average infant gains its length at the rate of 1.61cm per month with significant variability among infants (sd = 0.15; 95% CI: 0.12 to 0.18). For every infant the rate of gaining length over the following 16 month decreases at the rate of 0.03 cm per month (95% CI: -0.04 to -0.03). The result does not show significant correlation ($\hat{\rho} = -0.10$; 95%CI: -0.31 to 0.11) between initial length of infants and the rate that infant gains length over the following 16 months.

Table 7.1: Comparison of four alternative polynomial growth models fitted to length of infant data from the P-MaMiE study

		Model 1: No Change over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model 3 : Linear Change over time with random coefficients*	Model4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	66.87(66.7, 67.1)	58.84(58.6, 59.1)	58.85(58.65, 59.05)	57.56(57.32, 57.79)
	Age (linear term)		1.05(1.04, 1.07)	1.05(1.03, 1.07)	1.61(1.55, 1.66)
	Age^2(Quadratic term)				-0.03(-0.04, -0.03)
Variance component					
Level 1					
	Within person	6.79	3.20 (3.13, 3.28)	3.11(3.03, 3.20)	2.89(2.81, 2.97)
Level 2					
	Status at two months of age	1.16	2.24(2.09, 2.39)	1.96(1.72, 2.23)	2.16(1.94, 2.40)
	Linear term				
	Standard deviation			0.12(0.09, 0.17)	0.15(0.12, 0.18)
	correlation with at two month status			0.10(-0.25, 0.42)	-0.10(-0.31, 0.11)
Goodness of fit measures					
	Log likelihood	-14953.6	-12048.4	-12025.8	-11821.4
	AIC	29909.1	24104.9	24063.6	23656.8
	BIC	29915.5	24130.5	24102.0	23701.6

Figure 7.1 shows the predicted growth curve for selected infants implied by random coefficient quadratic function with observed growth data. For the majority of infants the random coefficients quadratic function fits to the observed data well.

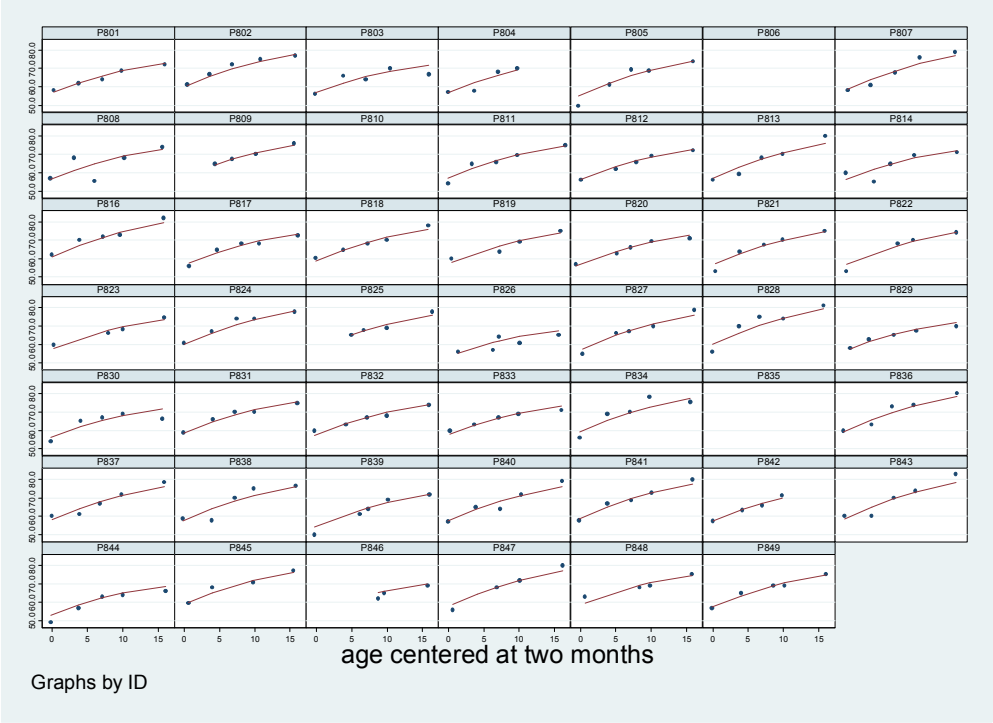

Figure 7.1: Trellis graph of observed length of selected infants (dots) and fitted trajectory using random coefficient (i.e. random intercept and random linear slope) quadratic growth model (dashed)

Histograms presented in figure 7.2 and 7.3 are approximately normal. Hence, the distributional assumption of normality for the three random components of quadratic mixed effects growth model fitted to the length of infants (i.e. random intercepts, random slopes and level-1 residuals) is not violated by the data.

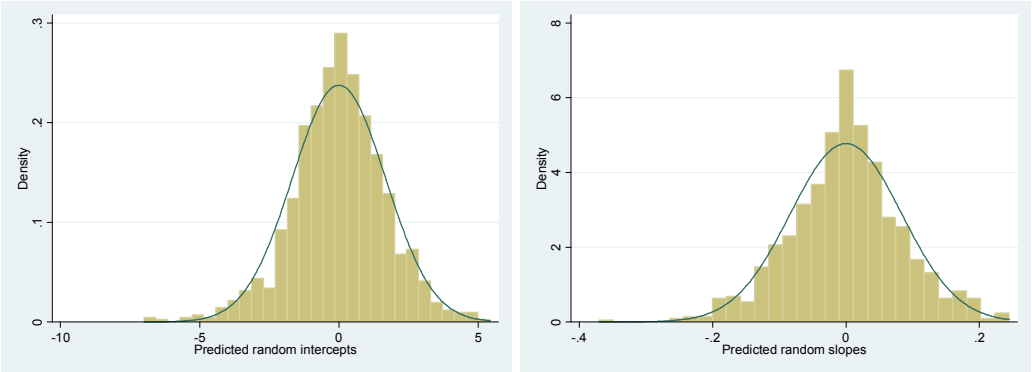

Figure 7.2: Histogram of predicted random slopes and random intercepts of the trajectory of length

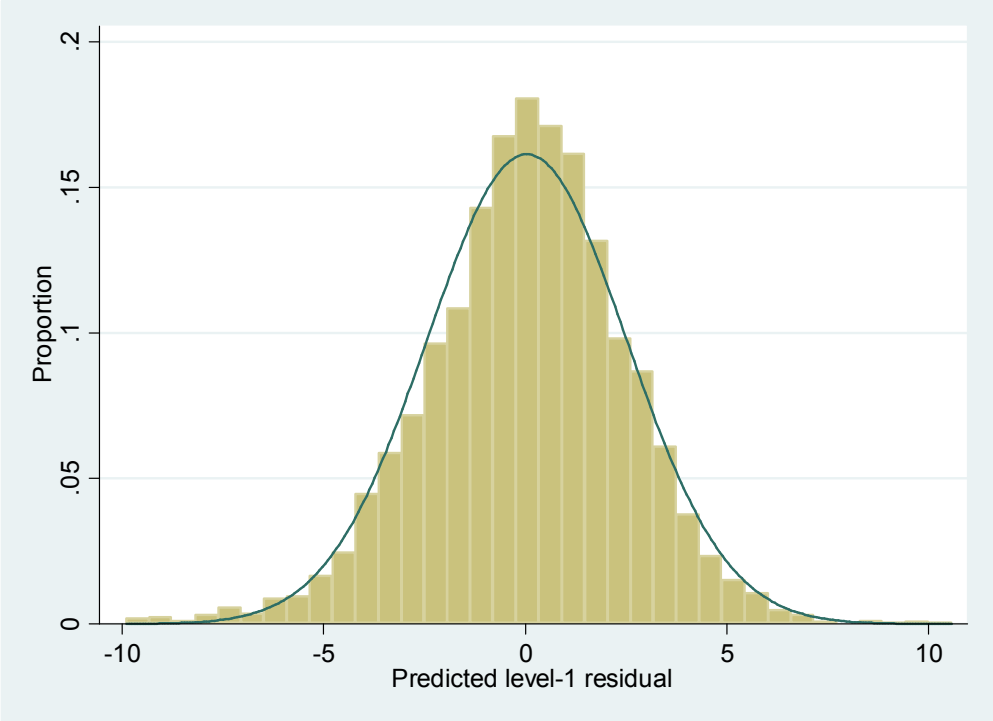

Figure 7.3: Histogram of predicted level-1(i.e. within infant) residual of the trajectory of length

7.2.2 Change in weight of infants over the first 18 months of age

Over the first 18 months of life the best fitting model among the candidate models evaluated in describing changes in weight of infants is a quadratic function with random intercepts and random slopes (table 7.2). There was improvement in model fit at every step of model comparison from "no change of weight over time but random intercepts" model to the "quadratic change in weight over time with random slopes and random intercepts" model. The values of Information Criteria were smaller, reduction in log-likelihood was statistically significant and within infant variability was reduced for every model that relaxes assumptions compared to the model which has relatively more restrictive assumptions. The available data was insufficient to enable us to estimate the standard errors of the random components of the quadratic component.

Summary results obtained from model 4 (table 7.2) shows that a two month old average infant attains mean weight of 5.18kg with significant variability between individual infants (sd = 0.47kg, 95% CI: 0.40 to 0.54). Starting at two months of age an average infant gains weight at a rate of 0.39kg/month with significant variability between infants (sd = 0.05; 95% CI: 0.04, 0.06). This rate of increase in weight decreases at a rate of 7.0 grams/month (95%CI: -8.0gms to -6.0gms)). Instantaneous rate of weight increase of an average infant at two months of age is positively correlated with the attained weight at this age ($\hat{\rho}$ = 0.74 ; 95% CI: 0.18 to 0.94).

Figure 7.4 shows fitted values of weight from the best fitting quadratic function with random coefficients superimposed on the scatter plot of observed weight for selected infants. Information summarized in this graph does not show significant deviations of observed weight measurements from the fitted values implying that quadratic function with random coefficients is enough to explain weight gain of infants of the P-MaMiE study over the first 18 months of life

The three histograms presented in figures 7.5 and 7.6 representing empirical distributions of the three random components of the best fitting quadratic function are normally distributed centred around the mean of zero. This shows that the assumption of normality for these random components underlying the fitted model is not severely violated

Table 7.2: Comparison of fitting alternative polynomial change trajectories to weight of infant data from the P-MaMiE study

		Model 1: No Chang over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model 3: Linear Change over time with random coefficients*	Model 4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	7.53(7.47, 7.59)	5.46(5.39, 5.53)	5.46(5.40, 5.51)	5.18(5.12, 5.24)
	Age (linear term)		0.27(0.27, 0.28)	0.27(0.27, 0.28)	0.39(0.38, 0.41)
	Age^2(Quadratic term)				-0.007 (-0.008, -0.006)
Variance component					
Level 1					
	Within person sd	1.87 (1.82, 1.91)	0.91(0.89, 0.93)	0.86(0.84, 0.89)	0.83(0.81, 0.85)
Level 2					
	Sd of status at two months of age	0.30(0.18, 0.47)	0.78(0.73, 0.82)	0.44(0.37, 0.52)	0.47(0.40, 0.54)
	Linear term				
	Standard deviation			0.05(0.04, 0.06)	0.05(0.04, 0.06)
	correlation with at two month status			0.92(-0.84, 1.00)	0.74(0.18, 0.94)
Goodness of fit measures					
	Log likelihood	-9215.3	-6630.4	-6489.3	-6368.3
	AIC	18436.6	13268.9	12990.7	12750.6
	BIC	18455.8	13294.5	13029.1	12795.5

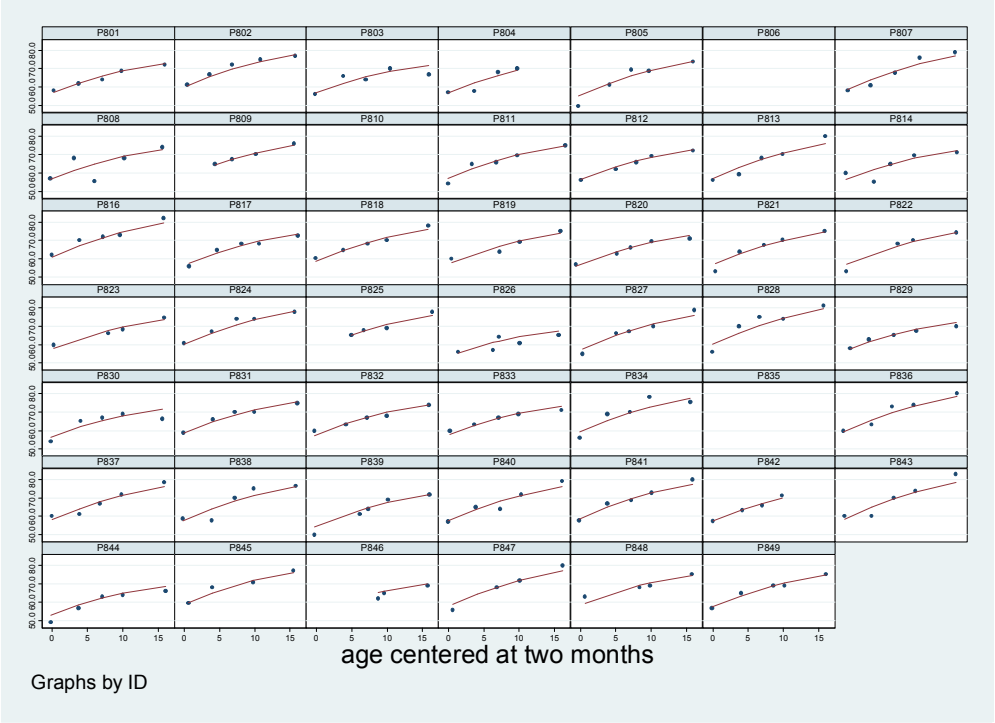

Figure 7.4: Trellis graph of observed weight of selected infants (dots) and fitted trajectory using random coefficient quadratic growth model (dashed)

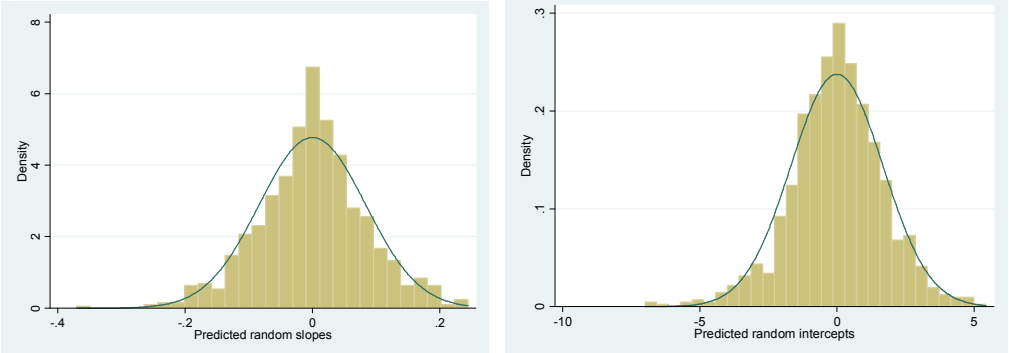

Figure 7.5: Histogram of predicted random slope and random intercept for the quadratic function of weight trajectory

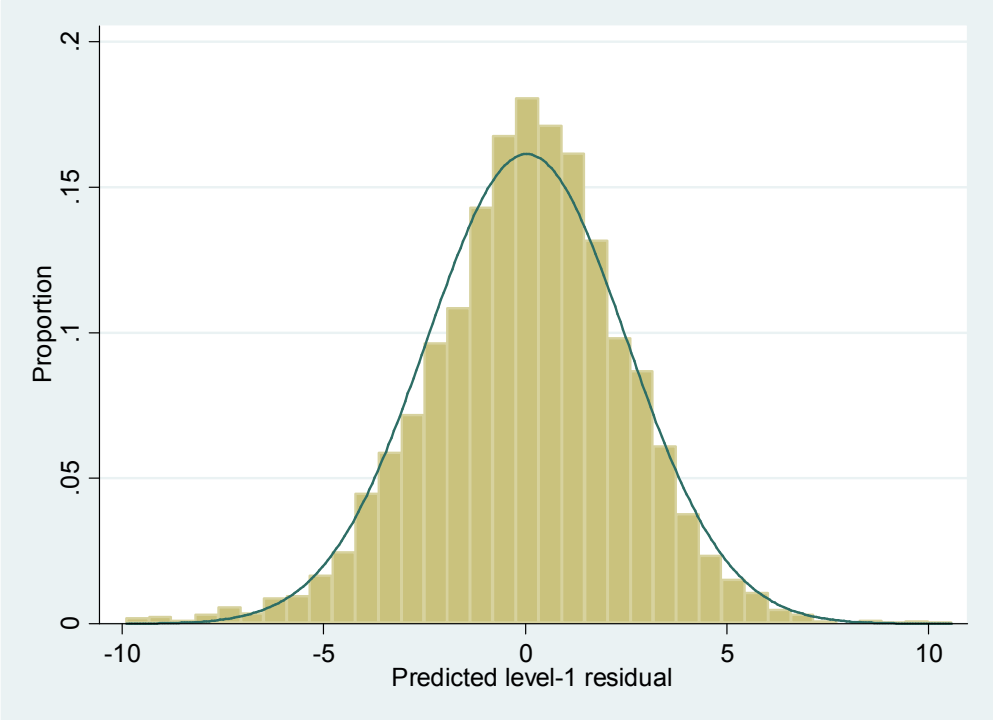

Figure 7.6: Histogram of predicted level-1(i.e. within infant) residual from quadratic function of weight trajectory

7.2.3 Changes in length-for-age z of infants over the first 18 months of age

Among four candidate models summarized in table 7.3 quadratic function with random intercepts and random slopes is the best fitting model in describing length-for-age z data of infants over the first 18 months of life. The available data was insufficient to enable us to estimate the standard errors of the random terms of the quadratic component. There was always improvement in model fit at every step of model comparison from a "no change of length-for-age z score with random intercept " model to the "quadratic change in length-for-age z score with random slopes and random intercepts" model. The values of Information Criteria were smaller, reduction in log-likelihood was statistically significant and within infant variability was reduced for every model in the hierarchy that relaxes assumptions compared to the model which has relatively more restrictive assumptions.

Table 7.3 Comparison of fitting alternative polynomial change trajectories to height-for-age z of infant data from the P-MaMiE study

		Model 1: No Change over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model3: Linear Change over time with random coefficients*	Model 4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	-1.52(-1.59, -1.45)	-0.53(-0.61, -0.44)	-0.53(-0.62, -0.44)	-0.23(-0.33, -0.13)
	Age (linear term)		-0.13(-0.14, -0.12)	-0.13(-0.14, -0.12)	-0.26(-0.28, -0.24)
	Age^2(Quadratic term)				0.008(0.006, 0.009)
Variance component					
Level 1					
	Within person sd	1.44(1.41, 1.48)	1.22(1.19, 1.25)	1.19(1.16, 1.23)	1.16(1.13, 1.20)
Level 2					
	sd of status at two months of age	0.88(0.82, 0.95)	0.93(0.87, 0.99)	0.98(0.89, 1.07)	0.98(0.90, 1.07)
	Linear term				
	standard deviation			0.04(0.03, 0.06)	0.05(0.03, 0.06)
	correlation with at two month status			-0.28(-0.48, -0.06)	-0.29(-0.47, -0.09)
Goodness of fit measures					
	Log likelihood	-8347.2	-7743.98	-7739.9	-7671.1
	AIC	16700.3	15496.0	15491.9	15356.2
	BIC	16719.5	15521.5	15530.2	15400.9

At the age of two months an average infant attains mean length-for-age z score of -0.23 (95% CI: -0.33, -0.13) with significant variability between individual infants (sd = 0.98, 95% CI: 0.90 to 1.07). Similarly, instantaneous rate of decrease in mean length-for-age z score of an average infant is 0.26 standard deviation units per month (95% CI: -0.28, -0.24) with significant variability between individual infants (sd = 0.05; 95% CI: 0.03, 0.06). The rate of decrease of length-for-age z score of an average infant (i.e. curvature of instantaneous decreases) also decreases at the rate of 0.008 standard deviation units per month (95% CI: 0.006, 0.009). At the age of two months instantaneous rate of decrease in length-for-age z score of an average infant is negatively correlated with the attained mean length-for-age z score ($\hat{\rho}$ = -0.29 ; 95% CI: -0.47 to -0.09) implying that infants who are well off at the age of two months are more likely to decrease their length-for-age z faster than those infants who had lower length-for-age z score**.**

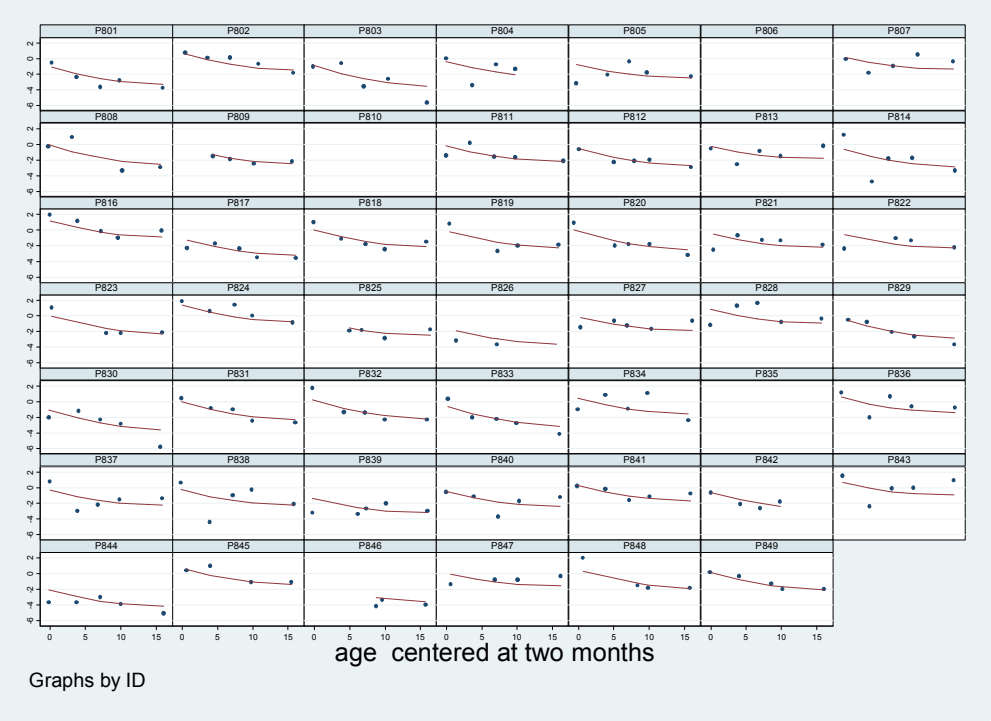

Figure 7.7: Trellis graph of observed height-for-age of selected infants (dots) and fitted trajectory using random coefficient quadratic growth model (dashed)

Figure 7.7 shows predicted values of length-for-age z score from the best fitting quadratic function superimposed on the scatter plot of observed length-for-age versus infant age for selected infants. This plot does not show significant deviations of observed z scores from the

predicted z score implying that quadratic function is enough to explain change in length-for-age z score of infants of the P-MaMiE study

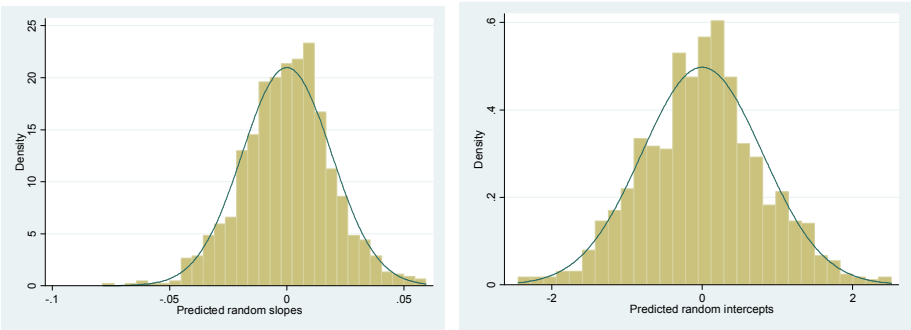

Figure 7.8: Histogram of predicted random slope and random intercept for the quadratic function of length-for-age z trajectory

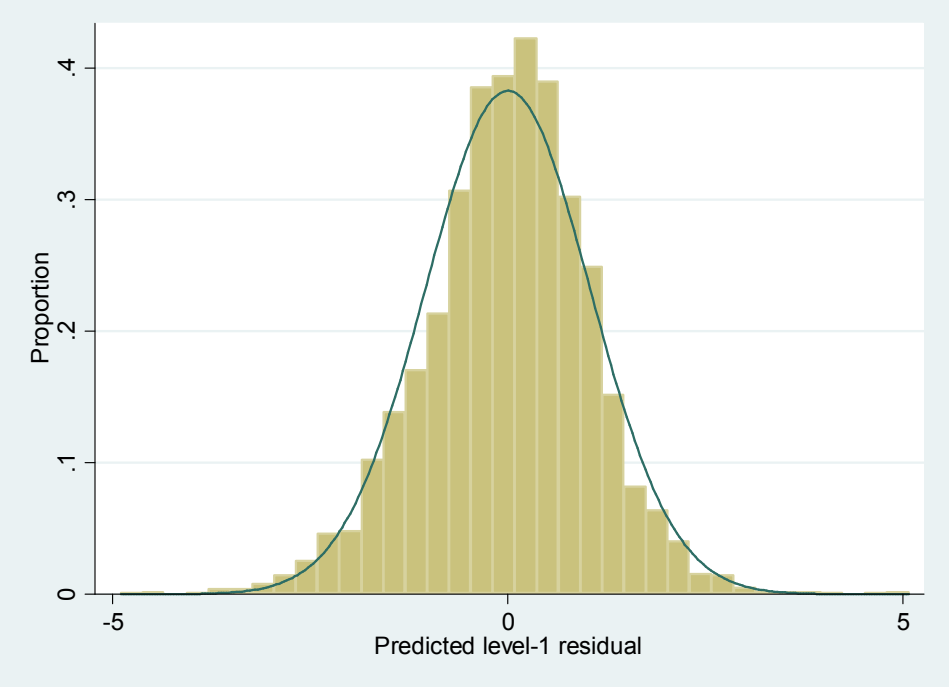

Figure 7.9: Histogram of predicted level-1(i.e. within infant) residual from quadratic function of length-for-age z trajectory

7.2.4 Changes in weight-for-age z score of infants over the first 18 months of life

Summary of results obtained from four models and presented in table 7.4 shows that the best fitting model to weight-for-age z data of infants over the first 18 months of life is quadratic function with random intercepts and random slopes. There was improvement in model fit at every step of model comparison from a "no change of weight-for-age z score with random intercept" model to the "quadratic change in weight-for-age z score with random slopes and random intercepts" model. The values of Information Criteria were smaller, reduction in log-likelihood was statistically significant and within infant variability over time was reduced for every model that relaxes assumptions compared to the model with relatively more restrictive assumptions.

At the age of two months an average infant attains mean weight-for-age z score of -0.55 (95% CI: -0.62 to -0.47) with significant variability between individual infants (sd = 0.81, 95% CI: 0.75 to 0.88). Instantaneous rate of decrease in mean weight-for-age z score of a two month old average infant is 0.14 standard deviation units per month (95% CI: -0.15 to - 0.12) and this rate of decline significantly vary between individual infants (sd = 0.04; 95% CI: 0.03, 0.05). The rate of decrease (i.e. curvature of instantaneous decreases) increases at the rate of 0.007 standard deviation units per month (95% CI: 0.006 to 0.008). At the age of two months instantaneous rate of decrease in weight-for-age z score is not significantly correlated with the attained weight-for-age z score. This implies that the rate at which weight-for-age z score of infants change over the first 18 months of life is not affected by the value attained at two months of age.

Figure 7.10 shows fitted values of weight-for-age z score from quadratic function with random coefficients model superimposed on the scatter plot of observed weight-for-age z versus age of infants for selected infants. The plot does not show significant deviations of observed z-scores from the fitted values implying that quadratic function with random coefficients is enough to explain the change in weight-for-age z score of infants of the P-MaMiE study. Histograms of the random components of the model presented in figures 7.11-7.12 also shows that the assumption of normality of these terms is not violated

Table 7.4: Comparison of fitting alternative polynomial change trajectories to weight-for-age of infant data from the P-MaMiE study

		Model 1: No Change over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model 3b: Linear Change over time with random coefficients*	Model 4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	-0.96(-1.02, -0.90)	-0.82(-0.89, -0.75)	-0.82(-0.89, -0.75)	-0.55(-0.62, -0.47)
	Age (linear term)		-0.02(-0.02, -0.01)	-0.02(-0.02, -0.01)	-0.14(-0.15, -0.12)
	Age^2(Quadratic term)				0.007(0.006, 0.008)
Variance component					
Level 1					
	Within person sd	0.96(0.94, 0.98)	0.95(0.93, 0.98)	0.94(0.91, 0.96)	0.90(0.88, 0.93)
Level 2					
	Standard deviation of				
	Status at two months	0.85(0.81, 0.90)	0.86(0.81, 0.91)	0.79(0.73, 0.87)	0.81(0.75, 0.88)
	Linear change			0.03(0.02, 0.05)	0.04(0.03, 0.05)
	Quadratic change				
	Correlation between				
	Linear terms and intercept			0.16(-0.20, 0.48)	0.04(-0.02, 0.28)
	Linear term and quadratic				
	Intercept and quadratic term				
Goodness of fit					
	Log likelihood	-6877.2	-6854.2	-6842.9	-6742.7
	AIC	13760.5	13716.5	13697.8	13499.4
	BIC	13779.7	13742.1	13736.2	13544.2

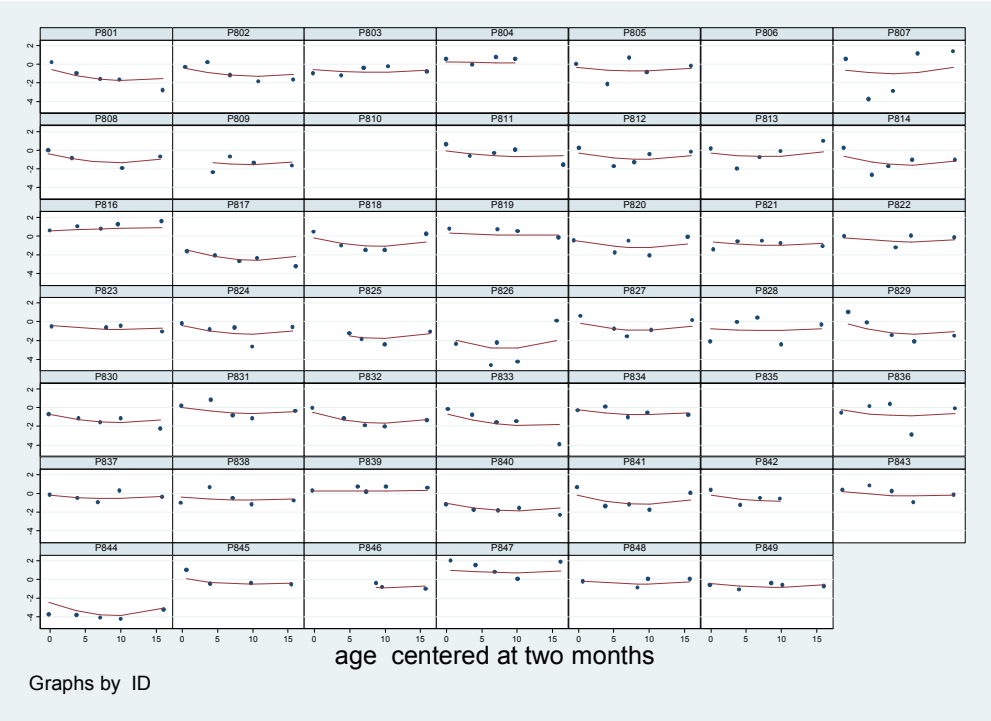

Figure 7.10: Trellis graph of observed weight-for-age z-score of selected infants (dots) and fitted trajectory using random coefficient quadratic growth model (dashed)

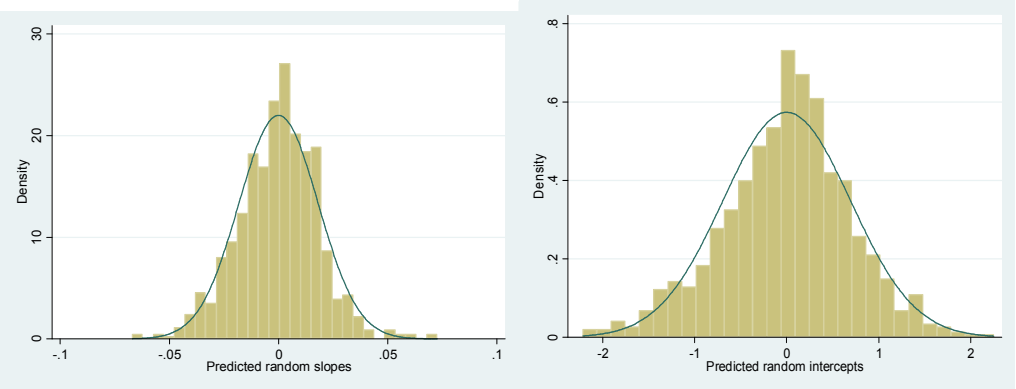

Figure 7.11: Histogram of random slopes and random intercepts of weight-for-age z

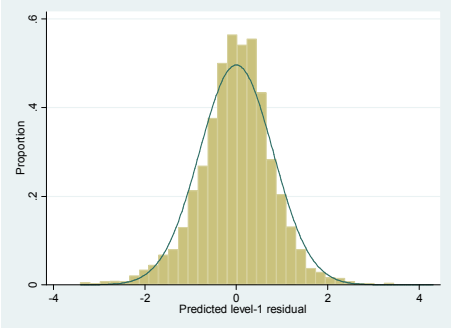

Figure 7.12 Histogram of level-1 residuals of weight-for-age z

7.2.5 **Changes in the logit of stunting of infants over the first 18 months of age**

The summary presented in table 7.5 shows that the best fitting model to the logit of stunting of infants over the first 18 months of life is linear with random intercepts and random slopes. In the hierarchy of fitting candidate models there was progressive improvement in model fit from a "no change in the logit of stunting with random intercept " model to the "linear change in the logit of stunting with random coefficients" model. However, inclusion of the quadratic term did not improve the fit of the model. For the first three models the values of information criteria were smaller, reduction in log-likelihood was statistically significant and variability of within individual growth was reduced for every model that relaxes assumptions compared to the model with relatively more restrictive assumptions. However, changes were not significant between model three and model four.

A two month old average infant attains mean value of the logit of stunting of -2.03 (95% CI: -2.23 to -1.83) with significant variability between individual infants (sd = 0.78, 95% CI: 0.47 to 1.30). From two to 18 months of age logit of stunting of an average infant decreases at a rate of 0.19/month (95% CI: 0.17, 0.22) with significant variability among individual infants (sd = 0.12; 95%CI: 0.09, 0.17). The rate of linear decreases of an average infant in the logit of stunting does not depend on the mean value of logit of stunting attained at the two months of age ($\hat{\rho}$ = 0.48 ; 95% CI: -0.54 to 0.93)

Table 7.5: Comparison of fitting alternative polynomial change trajectories to the logit of stunting of infant data from the P-MaMiE study

		Model 1: No Change over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model3: Linear Change over time with random coefficients*	Model 4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	-0.54(-0.64, -0.45)	-2.14(-2.33, -1.95)	-2.03(-2.23, -1.83)	-2.20(-2.49, -1.91)
	Age (linear term)		0.19(0.18, 0.21)	0.19(0.17, 0.22)	0.25(0.18, 0.31)
	Age^2(Quadratic term)				-0.003(-0.007, 0.001)
Variance component					
Level 1					
	Within person sd				
Level 2					
	Sd of status at two months of age	1.01(0.90, 1.14)	1.39(1.25, 1.54)	0.78(0.47, 1.30)	0.93(0.60, 1.43)
	Linear term				
	Standard deviation			0.12(0.09, 0.17)	0.12(0.08, 0.17)
	correlation with at two month status			0.48(-0.54, 0.93)	0.28(-0.44, 0.78)
Goodness of fit					
	Log likelihood	-2853.9	-2505.5	-2479.1	-2477.7
	AIC	5711.9	5017.0	4968.1	4967.3
	BIC	5724.7	5036.2	5000.1	5005.7

7.2.6. Changes in the logit of underweight of infants over the first 18 months of age

Of the four nested models evaluated with the data of the logit of underweight of infants and summarized in table 7.6 quadratic function with random intercepts and random slopes is the best fitting model. Although there is progressive improvement in model fit from a "no change in the logit of underweight with random intercept" model to the "linear change in the logit of underweight with random coefficients" model, improvement of model fit gained at each step was not large. However, inclusion of the quadratic term improved the fit significantly. The reduction in log-likelihood and an increase in each of the two Information Criteria are considerably large.

A two month old average infant attains a mean value of the logit of underweight of -2.83 (95% CI: -3.23 to -2.43) with significant variability between individual infants (sd = 1.49, 95% CI: 1.13 to 1.98). At this age instantaneous decrease of the logit of underweight of an average infant is 0.26 (95%CI: 0.18, 0.34) and it significantly vary across infants (sd = 0.15, 95% CI: 0.11 to 0.21). There is significant curvature ($\hat{\beta}$ = -0.02 ; 95% CI: -0.02 to -0.01) that accompanies the mentioned instantaneous rate of decrease. Instantaneous rate of change and mean value of the logit of underweight at two months of age are not significantly correlated.

Table 7.6 Comparison of fitting alternative polynomial change trajectories to logit of underweight of infant data from the P-MaMiE study

		Model 1: No Change over time with random intercept	*Model 2: Linear Change over time with random intercept*	*Model3: Linear Change over time with random coefficients*	Model 4: Quadratic change over time with random coefficients
		Estimate (95% CI)	*Estimate (95% CI)*	*Estimate (95% CI)*	Estimate (95% CI)
Fixed Effects					
	Intercept (status at two months)	-2.16(-2.35, -1.97)	-2.44(-2.67, -2.20)	-2.03(-2.28, -1.79)	-2.83(-3.23, -2.43)
	Age (linear term)		0.03(0.02, 0.05)	-0.02(-0.05, 0.01)	0.26(0.18, 0.34)
	Age^2(Quadratic term)				-0.02(-0.02, -0.01)
Variance component					
Level 1					
	Within person sd				
Level 2					
	Sd of status at two months of age	1.82(1.63, 2.03)	1.82(1.64, 2.04)	1.20(0.95, 1.52)	1.49(1.13, 1.98)
	Linear term				
	Standard deviation			0.09(0.06, 0.14)	0.15(0.11, 0.21)
	correlation with at two month status			1.00(-1.00, 1.00)	0.25(-0.24, 0.64)
Goodness of fit					
	Log likelihood	-1966.6	-1958.3	-1946.5	-1906.9
	AIC	3937.3	3922.5	3903.1	3825.9
	BIC	3950.1	3941.8	3935.1	3864.3

7.2.7 Growth levels attained by an average infant in the P-MaMiE study at selected ages of infancy

Growth levels attained by an average infant at the ages of 2, 6, 9, 12 and 18 months are presented in table 7.7. From two to eighteen months of age an average infant gains length and weight progressively. At the age of two months it attained 57.6 cm (95%CI: 57.3 to 57.8) in length, 5.2kg (95%CI: 5.1 to 5.2) in weight and it reached at a mean length of 74.7cm (95%CI: 74.7 to 74.9) and mean weight of 9.6kg (95%CI: 9.6 to 9.7) at the age of 18 months. This gain is equivalent to 30% increase in length and 84.6% increase in weight over 16 months of follow-up compared to the growth measurements attained at the two months of age. The increase in length and weight was monotonic. However, there was significant variability among infants at each age considered for the follow-up and this variability increased with the age of infants.

When standard deviation units (i.e. z-score) are used as growth measurement, an average infant in this cohort was shorter and lighter at two months of age (i.e. the starting point) compared to the expected value of zero in the 2006 WHO reference population. At two months of age an average infant in this cohort has a mean length-for-age z of -0.23 (95%CI: -0.33, -0.13) and a mean weight-for-age z of -0.55(95%CI: -0.62 to -0.47) which are both significantly smaller than zero. The weight-for-age z and length-for-age z of an average infant then continue to deteriorate as age increases and attains mean length-for-age z of -2.37(95%CI: -2.47 to -2.27) at the age of 18 months and mean weight-for-age z of -1.19(95%CI: -1.26 to -1.12) at the age of 12 months which are both significantly smaller than zero. Mean weight-for-age z at eighteen months of age is -0.89 (95%CI: -0.97 to -0.80) which is still significantly smaller than zero but better than it was at six months of age. Over the whole range of age studied there is significant variability on the attained length-for-age z and weight-for-age z. This variability among individual infants increases over time

Table 7.7: Attained population mean growth measures of infants at 2, 6, 9, 12 and 18 months of age in the P-MaMiE study

		Length	*weight*	*Length-for-age*	*Weight-for-age*	*Logit of stunting*	*Logit of underweight*
Fixed Effects							
	Two months	57.6(57.3, 57.8)	5.2(5.1, 5.2)	-0.23(-0.33, -0.13)	-0.55(-0.62,-0.47)	-2.03(-2.23,-1.83)	-2.83(-3.23,-2.43)
	Six months	63.5(63.3, 63.6)	6.6(6.6, 6.7)	-1.14(-1.21,-1.06)	-0.98(-0.04,-0.92)	-1.25(-1.40,-1.10)	-2.07(-2.30,-1.84)
	Nine months	67.2(67.0, 67.3)	7.6(7.5, 7.6)	-1.66(-1.73,-1.58)	-1.15(-1.22,-1.08)	-0.66(-0.80,-0.53)	-1.88(-2.10,-1.66)
	Twelve months	70.3(70.1, 70.5)	8.4(8.3, 8.4)	-2.04(-2.11,-1.96)	-1.19(-1.26,-1.12)	-0.08(-0.23, 0.07)	-2.02(-2.28,-1.75)
	Eighteen months	74.7(74.4, 74.9)	9.6(9.5, 9.7)	-2.37(-2.47, -2.27)	-0.89(-0.97,-0.80)	1.09(0.85, 1.33)	-3.28(-3.81,-2.75)
Random Part (sd of the attained growth level							
	Two months	2.16(1.94, 2.40)	0.47(0.40,0.54)	0.98(0.90, 1.07)	0.81(0.75,0.88)	0.78(0.47,1.30)	1.49(1.13, 1.98)
	Six months	2.18(2.03,2.34)	0.63(0.59,0.68)	0.94(0.88, 1.01)	0.83(0.78,0.88)	1.11(0.93, 1.31)	1.74(1.50, 2.02)
	Nine months	2.30(2.26,2.44)	0.77(0.72,0.81)	0.94(0.88, 1.00)	0.86(0.81,0.91)	1.42(1.26, 1.59)	2.01(1.79,2.26)
	Twelve months	2.49(2.34,2.65)	0.91(0.86,0.96)	0.96(0.90, 1.02)	0.90(0.85,0.96)	1.75(1.55, 1.98)	2.34(2.06, 2.67)
	Eighteen months	2.85(2.60,3.12)	1.20(1.13,1.27)	1.04(0.95, 1.14)	1.02(0.95,1.10)	2.45(2.09, 2.87)	3.13(2.62,3.73)
95% CI for attained growth by individual infants at a given time point							
	Two months	(53.4, 61.8)	(4.3, 6.1)	(-2.15, 1.69)	(-2.14, 1.04)	(-3.56, -0.50)	(-5.75, 0.09)
	Six months	(59.2, 67.8)	(5.1, 7.5)	(-2.98, 0.70)	(-2.61, 0.65)	(-3.43, 0.93)	(-5.48, 1.34)
	Nine months	(62.7, 71.7)	(6.1, 9.1)	(-3.50, 0.18)	(-2.84, 0.54)	(-3.44, 2.12)	(-5.82, 2.06)
	Twelve months	(65.4, 75.2)	(6.6, 10.2)	(-3.92, -0.16)	(-2.95, 0.57)	(-3.31, 3.35)	(-6.61, 2.57)
	Eighteen months	(69.1, 80.3)	(7.2, 12.0)	(-4.41, -0.33)	(-2.89, 1.11)	(-3.71, 5.89)	(-9.41, 2.85)

Within the cohort an average infant was not stunted in the first year of infancy and not underweight in the whole study period. Compared to the 2006 WHO reference population (a) there was no significant risk of underweight for an average infant during the first 18 months of follow-up, (b) the risk of stunting of an average infant increased with age, the risk was marginally non-significant at one year of age (OR = 0.92 95% CI: 0.79 to 1.07) and it was significantly increased at the age of 18 months (OR =2.97; 95% CI: 2.34 to 3.78). There was significant variability among individual infants in terms of the risk of stunting and underweight. This variability was increasing over time.

7.2.8 Summary

Four candidate polynomial functions were evaluated (i.e. random intercept but no growth over time, random intercept and non-random linear growth over time, random intercept and linear growth over time with random slopes, and quadratic growth over time with random intercepts and random slopes) to establish the best fitting function which describes growth of infants (i.e. length, weight, length-for-age z, weight-for-age z, logit of stunting and logit of underweight) between two and eighteen months of age. The time of growth measurements were centred at two months to facilitate the interpretation of the intercept term. With the exception of logit of stunting which was best described by linear growth with random intercepts and random slopes the best fitting function for all growth outcome measures was quadratic with random intercepts and random linear slopes.

A two month old average infant attained a length of 57.6 cm (95%CI: 57.3, 57.8), weight of 5.18kg(95%CI: 5.12, 5.24), length-for-age z of -0.23 (95%CI: -0.33, -0.13), weight-for-age z of -0.55(95%CI: -0.62, -0.47), logit of stunting of -2.20 (95%CI: -2.49, -1.91) and logit of underweight of -2.83(95%CI: -3.23, -2.43). There is significant variability between growth levels attained by two month old individual infants. All growth measurements of an average infant change significantly over time (i.e. an increase in length, weight, logit of stunting and logit of underweight, and a decrease in length-for-age z and weight-for-age z) with significant variability between individual infants. Initial length-for-age z and weight-for-age z are significantly and inversely correlated with linear slope of an average infant. However, the average infant's initial weight, length,

logit of stunting and logit of underweight are not significantly correlated with the linear slope.

Between two and eighteen months of age (a) there was a monotonic increase in weight and length of an average infant, (b) there was monotonic decline in length-for-age z and weight-for-age z compared to the 2006 WHO reference population and (c) there was no significant risk of underweight of an average infant. The risk of stunting of an average infant was statistically significant at the age of 18 months but not during the first year of infancy.

7.3. Crude and adjusted effects of different covariates on the initial status of infant growth

7.3.1 Length of an average infant at the age of two months

At the age of two months better maternal nutritional status, urban residence, lower score on poverty index, lower score on sanitary scale, male gender and being born with weight above 2500g were significant predictors of increased length for an average infant. A one centimetre increase in maternal MUAC was significantly associated with an increase of 0.11cm ($\hat{\beta}$ = 0.11; 95% CI: 0.03, 0.19) in the length of an average infant. An average infant living in an urban area was 1.29 cm longer than his rural counterpart ($\hat{\beta}$=1.29; 95% CI: 0.81, 1.76) and average female infant is 0.83cm shorter than an average boy ($\hat{\beta}$ = -0.83; 95% CI: -1.17, -0.50). An increase of one point in a poverty index and a similar increase in sanitary score are associated with a deficit of 0.13cm ($\hat{\beta}$ = -0.13; 95% CI: -0.22, -0.05) and a deficit of 0.29cm ($\hat{\beta}$ = -0.29, 95% CI: -0.47, -0.12), respectively. An average infant born with a weight of less than 2500g has a deficit of 1.65cm at the age of two months compared to a two month old average infant born with a weight of more than 2500g ($\hat{\beta}$=-1.65, 95% CI: -2.45, -0.85).

In the adjusted model colostrums become significant and length of a two month old an average infant who was not given colostrums is 0.57 cm (95% CI: 0.13, 1.02) longer than the one given colostrums. Other significant predictors of increased length for an average infant are better maternal nutritional status measured by MUAC ($\hat{\beta}$=0.12; 95%CI: 0.04, 0.20) and urban residence ($\hat{\beta}$= 1.69; 95%CI: 0.94, 2.43). Female gender ($\hat{\beta}$= -0.81; 95%CI: -1.15, -0.46) and low birth weight ($\hat{\beta}$= -1.78, 95%C: -2.60, -0.95) predicted shorted length for a two month old average infant compared to male infants and normal birth weight infants (BW >2500g), respectively.

7.3.2 Weight of an average infant at the age of two months

In an unadjusted model older parental age, maternal undernutrition, higher values of poverty index, increased values of sanitary scale, female gender, low birth weight and delayed initiation of breast feeding predicted lower weight of an average infant at the age

of two months. A one centimetre difference in maternal height and the same difference in maternal MUAC are associated with the difference of 10gm ($\hat{\beta}$=0.01; 95%CI: 0.001, 0.02) and 70gm ($\hat{\beta}$ = 0.07; 95% CI: 0.04, 0.09) in the average weight of a two month old infant, respectively, where the relationship is in the same direction. One unit increase in poverty index as well as on the sanitary scale is associated with decreases of weight of an average infant at the age of two months by 60gm ($\hat{\beta}$=-0.06; 95%CI: -0.09, -0.04) and 160gm ($\hat{\beta}$=-0.16; 95%CI: -0.21; 95%CI: -0.21, -0.11), respectively. Females are 290gm lighter than boys ($\hat{\beta}$=-0.29; 95%CI: -0.39, -0.20) and infants born with low birth weight are 720gm lighter ($\hat{\beta}$=-0.72; 95% CI: -0.95, -0.49) than infants born with normal birth weight. Infants who have started breastfeeding after an hour are 210gm lighter ($\hat{\beta}$= -0.21; 95% CI: -0.31, -0.10) than infants who have started breastfeeding within an hour of birth. Weight deficit of an average infant at the age of two months as the result of one year increase in father's age and one year increase in mother's age was 40gm and 10gm, respectively, which was marginally significant. Infants born from mothers with at least one obstetric complication are 160gm ($\hat{\beta}$=0.16; 95%CI: 0.06, 0.27) heavier than an average infant born from mother without any obstetric complication

In an adjusted model one centimetre increase in maternal MUAC is associated in an increase of 70gm ($\hat{\beta}$=0.07; 95%CI: 0.04, 0.09) in the weight of an average infant at two months of age. Infants born from mothers with at least one obstetric complication are 150gm ($\hat{\beta}$=0.15; 95%CI: 0.05, 0.25) heavier than an average infant born from mother without any obstetric complication. Average females infant is 310gm lighter ($\hat{\beta}$=-0.31; 95%CI: -0.41, -0.22) compared to an average male infant, low birth weight infant is 660gm lighter ($\hat{\beta}$=-0.66; 95%CI: -0.89, -0.43) compared to an average infant with normal birth weight, and infants who have started breastfeeding after one hour of birth are 150gm lighter ($\hat{\beta}$=-0.15; 95%CI: -0.25, -0.04) than an average infant who have started breastfeeding within an hour. One unit increase in sanitary scale is associated with a 100gm deficit ($\hat{\beta}$=-0.10; 95%CI: -0.16, -0.04) of an average infant's weight at the age of two months. Similarly, a one year increases in maternal age is associated with a deficit of a 10gm ($\hat{\beta}$=-0.01; 95%CI: -0.02, -0.00) of a two month old average infant's weight.

Table 7.8: Predictors of infant growth at initial point (2 months of age) in Butajira Birth cohort, Ethiopia

	Length		*Weight*	
	Unadjusted $\hat{\beta}$ (95% CI)	**Adjusted $\hat{\beta}$ (95% CI)**	**Unadjusted $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**
Characteristics of mother				
Age (years)	-0.01(-0.04, 0.02)	-0.03(-0.07, 0.01)	-0.01(-0.02, -0.001)	-0.01(-0.02, -0.00)
Height (metres)	0.01(-0.01, 0.04))	0.01(-0.02, 0.04)	0.01(0.001, 0.02)	0.01(-0.01, 0.01)
Mid upper arm circumference (cm)	0.11(0.03, 0.19)	0.12(0.04, 0.20)	0.07(0.04, 0.09)	0.07(0.04, 0.09)
Being in polygamous marriage	-0.08(-0.52, 0.35)	-0.22(-0.74, 0.30)	0.02(-0.11, 0.15)	0.07(-0.08, 0.21)
Autonomy scale (0-5)	-0.02(-0.11, 0.08)	-0.08(-0.18, 0.01)	0.01(-0.02, 0.04)	-0.01(-0.04, 0.02)
Use khat and/or alcohol	0.14(-0.29, 0.67)	0.15(-0.39, 0.68)	0.12(-0.03, 0.28)	0.13(-0.02, 0.28)
Had at least one obstetric complication	-0.09(-0.44, 0.28)	-0.05(-0.41, 0.31)	0.16(0.06, 0.27)	0.15(0.05, 0.25)
Household characteristics				
Urban residence	1.29(0.81, 1.76)	1.69(0.94, 2.43)	0.35(0.21, 0.49)	0.15(-0.06, 0.35)
Number of under 5 children: 0	-	-		-
1	-0.30(-0.74, 0.15)	-0.27(-0.75, 0.21)	-0.001(-0.13, 0.13)	0.03(-0.10, 0.17)
≥ 2	-0.41(-0.91, 0.08)	-0.49(-1.03, 0.04)	-0.04(-0.18, 0.11)	-0.08(-0.23, 0.07)
Age of father in years	-0.002(-0.02, 0.02)	0.01(-0.02, 0.04)	-0.01(-0.01, -0.00)	-0.002(-0.01, 0.01)
Poverty index (0 - 11)	-0.13(-0.22, -0.05)	0.05(-0.07, 0.18)	-0.06(-0.09, -0.04)	-0.01(-0.04, 0.03)
Poor sanitary condition (0 -3)	-0.29(-0.47, -0.12)	-0.10(-0.31, 0.12)	-0.16(-0.21, -0.11)	-0.10(-0.16, -0.04)
Level of social support (0 - 4)	-0.03(-0.18, 0.13)	0.02(-0.14, 0.18)	-0.02(-0.06, 0.03)	0.001(-0.04, 0.05)
Characteristics of index child				
Female gender	-0.83(-1.17, -0.50)	-0.81(-1.15, -0.46)	-0.29(-0.39, -0.20)	-0.31(-0.41, -0.22)
Not immunised at two months	0.27(-0.07, 0.61)	-0.34(-0.01,0.70)	-0.01(-0.11, 0.09)	0.02(-0.08, 0.12)
Severe illness in the first 2 months	-0.002(-0.42, 0.41)	-0.03(-0.46, 0.39)	-0.04(-0.16, 0.08)	-0.05(-0.17, 0.07)
Birth weight: Normal (>=2500g)	-	-	-	-
Low (<2500g)	-1.65(-2.45, -0.85)	-1.78(-2.60, -0.95)	-0.72(-0.95, -0.49)	-0.66(-0.89, -0.43)
Not measured	-0.34(-0.70, 0.02)	-0.49(-0.86, -0.11)	0.03(-0.07, 0.14)	0.01(-0.10, 0.11)
Early Infant Feeding practices				
Non-exclusive breast-feeding at 2 months	-0.06(-0.52, 0.40)	-0.11(-0.58, 0.36)	0.05(-0.09, 0.18)	0.03(-0.10, 0.16)
No pre-lacteal food	0.42(-0.38, 1.23)	0.65(-0.15, 1.45)	0.01(-0.22, 0.25)	0.04(-0.18, 0.27)
Colostrums not given	0.41(-0.03, 0.84)	0.57(0.13, 1.02)	-0.10(-0.23, 0.02)	0.01(-0.11, 0.14)
Breast feeding delayed for 1 hour	-0.21(-0.57, 0.16)	0.00(-0.37, 0.37)	-0.21(-0.31, -0.10)	-0.15(-0.25, -0.04)

7.3.3 Length-for-age z of an average infant at the age two months

Predictors of length-for-age z of a two month old average infant are summarized in table 7.9. In a bivariate analysis increased maternal MUAC, urban residence, female gender, and not receiving colostrums are significantly associated with increased length-for-age z score while having one previous child aged less than five years compared to no such child, increased value of poverty scale, increased value of poor sanitary scale and being born with low birth weight are inversely associated with increased length-for-age z. An increase of one centimetre in maternal MUAC, being resident of rural area compared to rural area, being a female infant, and not receiving colostrums at birth are associated with their respective increase of 0.04 standard deviation units (95%CI: 0.01, 0.08), 0.53 standard deviation units (95%CI: 0.33, 0.72), 0.46 standard deviation units (95%CI: 0.32, 0.59) and 0.18 standard deviation units (95%CI: 0.003, 0.36). Similarly, having one previous sibling aged more than 5 years compared to no such sibling, a one unit increase in poor sanitary score, a one unit increase in poverty scale and being born with low birth weight are associated with their respective deficit of 0.19 standard deviation units (95%CI: -0.37, -0.01), 0.15 standard deviation units (95%CI: -0.22, -0.08), 0.05 standard deviation units (95%CI:-0.09, -0.02) and 0.55 standard deviation units(95%CI: -0.88, -0.22).

In the adjusted model one year increase in maternal age is associated with a marginal deficit of a two month old average infant's length-for-age z. A one centimetre increase in maternal MUAC, being resident of urban area, being female infant, being not given colostrums and not receiving pre-lacteal food are associated with respective increases of a two month old average infant's length-for-age z score by 0.05 standard deviation unit (95%CI: 0.01, 0.08), 0.72 standard deviation units (95%CI: 0.42, 1.02), 0.47 standard deviation units (95%CI: 0.33, 0.61), 0.25 standard deviation units (95%CI: 0.07, 0.43) and 0.37 standard deviation units (95%CI: 0.05, 0.73). A two month old average infant who was born with low birth weight attains length-for-age z score which is 0.58 standard deviation unit (95%CI: -0.91, -0.24) lower than an average infant who was born with normal birth weight.

7.3.4 Weight-for-age z of an average infant at the age of two months

In unadjusted conditional multilevel growth model several factors are significantly associated with weight-for-age z score of a two month old average infant. A one year increase of maternal age, a one year increase in the age of the father, a one unit increases in poverty index score, a one unit increases in the score of poor sanitary condition scale, being born with low birth weight compared to an average infant born with normal birth weight, not being given colostrums compared to being given colostrums and delayed initiation of breastfeeding for more than an hour versus within an hour of birth are all associated with reduced weight-for-age z score where their estimated reduction in standard deviation unit are -0.02 (95CI: -0.03, -0.01), -0.01(95%CI: -0.02, -0.003), -0.08 (95%CI: -0.11, -0.05), -0.23 (95%CI: -0.29, -0.17), -0.96 (95%CI: -1.25, -0.68), -0.17 (95%CI: -0.33, -0.01), and -0.23 (95%CI: -0.37, -0.10), respectively. On the other hand a one centimetre increased in material MUAC, having mother with at least one obstetric complication compared with the one whose mother had no obstetric complication, being urban resident than rural resident, and being a female than male infant are all significantly associated with increased values of weight-for-age z score where estimated increase in standard deviation units are 0.08 (95%C: 0.05, 0.11), 0.22 (95%CI: 0.09, 0.35), 0.43 (95%CI: 0.23, 0.60), and 0.34 (95%CI: 0.22, 0.46), respectively.

After adjusting for other correlates a one centimetre increase in maternal MAUC, having a mother with at least one obstetric complication compared to no such complication, and being a female than being a male infant are all significantly associated with an increased weight-for-age z score with estimated increase (in standard deviation units) are 0.08 (95%CI: 0.05, 0.10), 0.17 (95%CI: 0.05, 0.30) and 0.31(95%CI: 0.19, 0.43), respectively. A one year increase of the mother's age, a one unit increase in the poor sanitary condition score, being born with low birth weight than with normal birth weight, and delayed initiation of breastfeeding for more than an hour than within an hour are all significantly associated with a decrease of weight-for-age z score where their associated decline (in standard deviation units) being -0.02 (95%CI: -0.03, -0.004), -0.15 (95%CI: -0.23, -0.08), -0.87(95%CI: -1.15, -0.58), and -0.15 (95%CI: -0.27, -0.02), respectively.

Table 7.9: Predictors of infant growth at initial point (2 months of age) in Butajira Birth cohort, Ethiopia

	Length-for-age		*Weight-for-age*	
	Unadjusted $\hat{\beta}$ (95% CI)	**Adjusted $\hat{\beta}$ (95% CI)**	**Unadjusted $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**
Characteristics of mother				
Age (years)	-0.01(-0.02, 0.00)	-0.02(-0.03, -0.00)	-0.02(-0.03, -0.01)	-0.02(-0.03, -0.004)
Height (metres)	0.003(-0.01, 0.01))	0.002(-0.01, 0.01)	0.01(-0.002, 0.02)	0.01(-0.002, 0.02)
Mid upper arm circumference (cm)	0.04(0.01, 0.08)	0.05(0.01, 0.08)	0.08(0.05, 0.11)	0.08(0.05, 0.10)
Being in polygamous marriage	0.05(-0.14, 0.23)	-0.01(-0.22, 0.19)	0.01(-0.15, 0.17)	0.05(-0.13, 0.23)
Autonomy scale (0-5)	-0.01(-0.04, 0.03)	-0.03(-0.07, 0.01)	0.01(-0.02, 0.04)	-0.01(-0.05, 0.02)
Use khat and/or alcohol	0.09(-0.12, 0.31)	0.07(-0.14, 0.29)	0.15(-0.04, 0.34)	0.14(-0.05, 0.32)
Had at least one obstetric complication	-0.02(-0.16, 0.13)	-0.03(-0.18, 0.11)	0.22(0.09, 0.35)	0.17(0.05, 0.30)
Household characteristics				
Urban residence	0.53(0.33, 0.72)	0.72(0.42, 1.02)	0.43(0.26, 0.60)	0.16(-0.10, 0.41)
Number of under 5 children: 0	-	-		-
1	-0.19(-0.37, -0.01)	-0.08(-0.27, 0.12)	-0.14(-0.30, 0.02)	-0.003(-0.17, 0.16)
≥ 2	-0.13(-0.33, 0.07)	-0.16(-0.38, 0.05)	-0.08(-0.25, 0.10)	-0.11(-0.29, 0.08)
Age of father in years	-0.002(-0.01, 0.01)	0.004(-0.01, 0.02)	-0.01(-0.02, -0.003)	-0.001(-0.01, 0.01)
Poverty index (0 - 11)	-0.05(-0.09, -0.02)	0.03(-0.02, 0.08)	-0.08(-0.11, -0.05)	-0.01(-0.05, 0.04)
Poor sanitary condition (0 -3)	-0.15(-0.22, -0.08)	-0.08(-0.16, 0.01)	-0.23(-0.29, -0.17)	-0.15(-0.23, -0.08)
Level of social support (0 - 4)	-0.03(-0.09, 0.04)	-0.02(-0.09, 0.04)	-0.03(-0.08, 0.03)	-0.01(-0.06, 0.05)
Characteristics of index child				
Female gender	0.46(0.32, 0.59)	0.47(0.33, 0.61)	0.34(0.22, 0.46)	0.31(0.19, 0.43)
Not immunised at two months	0.09(-0.05, 0.23)	0.13(-0.01, 0.27)	-0.002(-0.13, 0.12)	0.05(-0.07, 0.17)
Severe illness in the first 2 months	-0.07(-0.24, 0.10)	-0.01(-0.19, 0.16)	-0.11(-0.26, 0.04)	-0.07(-0.21, 0.08)
Birth weight: Normal (>=2500g)	-	-	-	-
Low (<2500g)	-0.55(-0.88, -0.22)	-0.58(-0.91, -0.24)	-0.96(-1.25, -0.68)	-0.87(-1.15, -0.58)
Not measured	-0.21(-0.36, -0.06)	-0.23(-0.38, -0.08)	-0.03(-0.16, 0.10)	-0.04(-0.17, 0.09)
Early Infant Feeding practices				
Non-exclusive breast-feeding at 2 months	-0.04(-0.23, 0.15)	-0.01(-0.20, 0.18)	0.01(-0.15, 0.18)	0.03(-0.13, 0.19)
No pre-lacteal food	0.28(-0.04, 0.61)	0.37(0.05, 0.73)	0.03(-0.26, 0.32)	0.08(-0.19, 0.26)
Colostrums not given	0.18(0.003, 0.36)	0.25(0.07, 0.43)	-0.17(-0.33, -0.01)	-0.001(-0.15, 0.15)
Breast feeding delayed for 1 hour	-0.05(-0.20, 0.10)	0.05(-0.10, 0.20)	-0.23(-0.37, -0.10)	-0.15(-0.27, -0.02)

7.3.5 Stunting of an average infant at the age of two months

In unadjusted conditional multilevel growth model one centimetre increase in maternal MUAC, being born from urbanely residing parents, and being a female infant are protective for an average infant aged two months from becoming stunted with their respective odds ratio (95%CI) of 0.93(0.88, 0.99), 0.48(0.33, 0.70) and 0.42(0.33, 0.53). In the same analysis an average infant aged two months is at an increased risk of being stunted if (a) it has more than one under five year sibling (OR = 1.13, 95%CI: 1.25, 1.60), (b) the family scored higher on poverty scale (OR = 1.07; 95%CI: 1.01, 1.13), (c) the family scored higher on a poor sanitary index (OR= 1.17; 95%CI: 1.03, 1.34), (d) it is born with low birth weight (OR = 1.93; 95%CI: 1.12, 3.32) and, (e) it is non exclusive breastfed before the age of two month(OR = 1.04; 95%CI: 1.32, 1.43).

In the adjusted model significant protective factors against stunting for a two month old average infant are an increased maternal MUAC (OR = 0.92; 95%CI: 0.88, 0.98), urban residence (OR = 0.30; 95%CI: 0.18, 0.51), female sex (OR = 0.42; 95%CI: 0.33, 0.53) and not being immunized (OR = 0.79; 95% CI: 0.62,0.99). The only significant factor that increased the risk of stunting in this study population is low birth weight (OR = 1.95; 95%CI: 1.14, 3.39).

7.3.6 Underweight of an average infant at the age of two months

Before adjusting for other risk factors higher score on poverty index, higher score on poor sanitary scale, low birth weight and delayed initiation of breast feeding were directly related with an increased value of underweight while increased value of maternal MUAC, urban residence, being born from a mother with at least one obstetric complication and being a female infant are inversely related with underweight of an average infant at the age of two months. A one centimetre increase in maternal MUAC, having a mother with at least one obstetric complication compared to infants whose mother has no obstetric complication, being a female compared to male infant, and being resident of urban area than being rural resident is associated with significantly reduced risk of being underweight with their respective odds ration of OR = 0.85(95%CI:0.79, 0.92), OR = 0.64 (95%CI: 0.45, 0.90), OR = 0.41(95%CI:0.29, 0.57) and OR = 0.41 (95%CI: 0.24, 0.68). A one point increase in poverty index score, a one point increase

in poor sanitary condition score, being born with low birth weight and delayed initiation of breast feeding for more than one hour increases the risk of underweight with respective odds rations of 1.14 (95%CI: 1.04, 1.25), 1.65 (95%CI: 1.38, 1.97), 7.24 (95%CI: 3.56, 14.59), and 1.65 (1.14, 2.39).

In an adjusted conditional multilevel growth model a one centimetre increase of maternal MUAC (OR = 0.84 95%CI: 0.78, 0.92) and being a female infant (OR = 0.44; 95%CI: 0.32, 0.62) have a protective effect against being underweight for a two month old average infant. However, a one point increase in poor sanitary condition score (OR = 1.49; 95%CI: 1.22, 1.82), low birth weight (OR= 6.05; 95%CI: 2.97, 12.30) and delayed initiation of breastfeeding for more than an hour (OR= 1.45; 95%CI: 1.01, 2.10) are significantly associated with an increased risk of being underweight.

Table 7.10: Predictors of infant undernutrition at initial point (2 months of age) in Butajira Birth cohort, Ethiopia

	Stunting		*Underweight*	
	Unadjusted $\hat{\beta}$ (95% CI)	**Adjusted $\hat{\beta}$ (95% CI)**	**Unadjusted $\hat{\beta}$ (95% CI)**	**Adjusted $\hat{\beta}$ (95% CI)**
Characteristics of mother				
Age (years)	0.01(-0.01, 0.03)	0.02(-0.01, 0.04)	0.03(-0.002, 0.05)	0.03(-0.01, 0.07)
Height (metres)	-0.004(-0.02, 0.02))	-0.006(-0.02, 0.01)	-0.02(-0.05, 0.005)	-0.02(-0.05, 0.01)
Mid upper arm circumference (cm)	-0.07(-0.13, -0.01)	-0.08(-0.13, -0.02)	-0.16(-0.24, -0.08)	-0.17(-0.25, -0.08)
Being in polygamous marriage	-0.05(-0.36, 0.26)	0.01(-0.27, 0.42)	0.008(-0.41, 0.43))	-0.03(-0.52, 0.46)
Autonomy scale (0-5)	0.01(-0.05, 0.08)	0.03(-0.03, 0.10)	-0.03(-0.12, 0.06)	0.01(-0.08, 0.10)
Use khat and/or alcohol	-0.24(-0.62, 0.14)	-0.19(-0.55, 0.18)	-0.40(-0.93, 0.14)	-0.32(-0.85, 0.22)
Had at least one obstetric complication	-0.11(-0.36, 0.14)	-0.03(-0.27, 0.21)	-0.45(-0.79, -0.10)	-0.27(-0.61, 0.08)
Household characteristics				
Urban residence	-0.74(-1.11, -0.36)	-1.20(-1.71, -0.68)	-0.90(-1.43, -0.38)	-0.34(-0.10, 0.41)
Number of under 5 children: 0	-	-		-
1	0.31(-0.002, 0.62)	0.17(-0.16, 0.50)	0.01(-0.47, 0.49)	0.03(-0.45, 0.50)
≥ 2	0.12(0.22, 0.47)	0.14(-0.22, 0.50)	0.01(-0.47, 0.49)	0.27(-0.25, 0.79)
Age of father in years	0.004(-0.01, 0.02)	-0.005(-0.03, 0.02)	0.02(-0.002, 0.04)	0.01(-0.02, 0.04)
Poverty index (0 - 11)	0.07(0.01, 0.13)	-0.03(-0.11, 0.05)	0.13(0.04, 0.22)	-0.03(-0.15, 0.09)
Poor sanitary condition (0 -3)	0.16(0.03, 0.29)	0.04(-0.10, 0.18)	0.50(0.32, 0.68)	0.40(0.20, 0.60)
Level of social support (0 - 4)	0.01(-0.10, 0.13)	0.002(-0.11, 0.11)	0.06(-0.10, 0.22)	0.03(-0.13, 0.18)
Characteristics of index child				
Female gender	-0.86(-1.10, -0.63)	-0.87(-1.10, -0.64)	-0.90(-1.23, -0.57)	-0.81(-1.15, -0.48)
Not immunised at two months	-0.22(-0.46, 0.02)	-0.24(-0.48, -0.01)	-0.09(-0.43, 0.24)	-0.15(-0.49, 0.19)
Severe illness in the first 2 months	0.11(-0.18, 0.40)	-0.07(-0.35, 0.22)	0.28(-0.11, 0.68)	0.20(-0.20, 0.60)
Birth weight: Normal (>=2500g)	-	-	-	-
Low (<2500g)	0.66(0.11, 1.20)	0.67(0.13, 1.22)	1.98(1.27, 2.68)	1.80(1.09, 2.51)
Not measured	0.36(0.11, 0.61)	0.39(0.14, 0.64)	-0.01(-0.37, 0.34)	-0.04(-0.41, 0.32)
Early Infant Feeding practices				
Non-exclusive breast-feeding at 2 months	0.04(0.28, 0.36)	-0.02(-0.33, 0.29)	0.01(-0.44, 0.45)	-0.09(-0.54, 0.35)
No pre-lacteal food	-0.20(-0.74, 0.35)	-0.31(-0.84, 0.22)	-0.05(-0.83, 0.73)	-0.14(-0.91, 0.64)
Colostrums not given	-0.14(-0.45, 0.17)	-0.20(-0.50, 0.10)	0.36(-0.06, 0.77)	-0.03(-0.45, 0.39)
Breast feeding delayed for 1 hour	0.11(-0.14, 0.37)	0.02(-0.23, 0.27)	0.50(0.13, 0.87)	0.37(0.01, 0.74)

7.3.7 Summary

Risk factors across all the six domains of growth

Before and after adjusting for other risk factors maternal MUAC, infant gender and birth weight were significantly associated with all six growth measures of a two month old infant. Higher maternal MUAC and low birth weight were significantly associated with better child growth and compromised child growth, respectively. Females compared to male infants weighed less, were shorter, scored better on weight-for-age z and length-for-age z, and were less likely to be stunted or underweight.

In unadjusted model urban residence, lower score on poverty index and lower score on poor sanitary condition scale were all significantly associated with better growth of infants on all growth measures at the age of two months. In a fully adjusted model urban residence was significantly associated with better attainment of the three length related growth measures and lower score on poor sanitary scale was significantly associated with better attainment of the three weight related growth measures. Maternal age was inversely related with length-for-age z only in a fully adjusted model and it was inversely associated with weight and weight-for-age z before and after adjusting for other risk factors. Infants who were not given colostrum scored relatively better on length-for-age z and scored relatively lower on weight-for-age z in unadjusted model. In a fully adjusted model, these infants were relatively taller and scored better on length-for-age z.

Risk factors specific to length related growth outcomes (length, length-for-age z and stunting)

Number of siblings less than five year of age, non-exclusive breast feeding before the age of two months and pre-lacteal feeding were significantly associated with one or more length related growth measures. In unadjusted model infants with one older sibling was more likely to score lower on length-for-age z compare to those who have no older sibling and infants having one or more siblings, or who were non-exclusively breastfed were more likely to be stunted. In a fully adjusted model those infants who were given pre-lacteal fed scored relatively lower on length-for-age z and infants unvaccinated before two months of age were less likely to be stunted.

Risk factors specific to weight related growth outcomes (weight, weight-for-age z and underweight)

Maternal height, obstetric complication, age of the father and delayed initiation of breast feeding were significantly associated with at least one of the weight, weight-for-age z and underweight of two month old average infant. At the age of two months delayed initiation of breast feeding for more than an hour was significantly associated with lighter weight, lower score on weight-for-age z and increased likelihood of underweight before and after adjusting for other risk factors. Similarly, increased father's age was significantly associated with compromised weight and weight-for-age z before adjusting for other risk factors but not in a fully adjusted model. In unadjusted model but not in a fully adjusted model increased maternal height was significantly associated with increased weight of a two month old average infant. Having a mother with one or more obstetric complication was significantly associated with reduced likelihood of underweight, an increased score on weight-for-age z and an increased weigh before adjusting for other risk factors. In a fully adjusted model the association with increased weight again and higher weight-for-age z score remained statistically significant.

7.4. Conditional MGM – Crude and adjusted effects of level-2 predictors of growth attainment at the age of two months and the rates at which growth outcomes change over time

All the tables in this section have the same format: results from three multilevel growth models are presented in each table. The first column describes the name of covariates with their corresponding sub-categories and the second column presents coefficients from a separate model for each covariate with main effect and its interaction with time. The third column presents change in deviance as a measure of model fit comparing the two nested models with only one covariate in the model, one model allowing interaction with time and the other model not. The last column of each table summarizes coefficients from a fully adjusted model in which all the covariates are allowed to predict the initial value (i.e. growth attainment a two months of age) and some of the covariates are allowed to predict the instantaneous rate of growth outcome (i.e. rate of growth over time).

7.4.1 Length of infants

Results from the following three multilevel growth models are presented in table 7.11: (a) the model with only one covariate which is allowed to predict the initial value and rate of gain in length in the second column of the table, (b) change in deviance as a measure of model fit comparing two nested change with only one covariate – one allowing interaction with time and the other not - and (c) the fully adjusted model in which all the covariates are allowed to predict the initial value and some of the covariates are also allowed to predict the instantaneous rate of increase in length in the first 18 months of life. In the unadjusted models seven variables showed significant interaction with time resulting in large reduction of deviance.

In unadjusted conditional multilevel growth models significant predictors of length attained at two months of age were urban residence ($\hat{\beta}$ =0.74; 95%CI: 0.16, 1.31), female gender of an infant ($\hat{\beta}$ = -0.81, 95%CI: -1.21, -0.41), and not getting colostrums ($\hat{\beta}$ = 0.52; 95%CI: 0.52; 95%CI: 0.01, 1.04). similarly, an increased maternal age ($\hat{\beta}$ = -0.01,

95% CI: -0.01, -0.002), increased maternal autonomy ($\hat{\beta}$ =0.01; 95%CI: 0.004, 0.02), urban residence of the family ($\hat{\beta}$ = 0.09; 95%CI: 0.04, 0.15), increased age of the father ($\hat{\beta}$ =-0.003; 95%CI: -0.01, -0.001), increased value on a poverty score ($\hat{\beta}$ = -0.01; 95% CI: -0.02, -0.01), and increased value of poor sanitary condition score ($\hat{\beta}$ = -0.07; 95%CI: -0.09, -0.05) were significant predictors of the instantaneous rate of change of infant length.

In a fully adjusted model maternal MUAC ($\hat{\beta}$ = 0.12; 95%CI: 0.03, 0.20), urban residence ($\hat{\beta}$ =1.67; 95%CI: 0.76, 2.58) , higher score on poor sanitary condition scale ($\hat{\beta}$ = 0.35; 95%CI: 0.09, 0.61) and not being given colostrums ($\hat{\beta}$ = 0.57; 95%CI: 0.12, 1.02) were positively associated with infant length at the age of two months while higher maternal autonomy score ($\hat{\beta}$ = -0.15, 95% CI: -0.27, -0.04), being female infant ($\hat{\beta}$ = -0.81; 95%CI: -1.15, -0.47), and being born with low birth weight ($\hat{\beta}$ =-0.1.75; 95%CI: -2.57, -0.92) were inversely associated with length of infant at two months of age. Similarly, an increase in maternal age ($\hat{\beta}$ = -0.01; 95%CI; -0.01, -0.001) and higher score on poor sanitary condition scale ($\hat{\beta}$ = -0.07; 95%CI: -0.09, -0.05) were inversely associated with instantaneous gain of length while higher score on maternal autonomy scale ($\hat{\beta}$ = 0.01; 95%CI: 0.00, 0.02) was positively associated with instantaneous rate of gaining length

Table 7.11: Interaction of different correlated of infant length at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ (95% CI)	Change in deviance	Adjusted $\hat{\beta}$ (95% CI)
Age (years)	0.02(-0.01,0.05)		-0.0004(-0.05,0.05)
Interaction with time	-0.01(-0.01,-0.002)	11.4	-0.005(-0.01, -0.00)
Height (metres)	0.002(-0.03,0.04)		0.01(-0.02,0.04)
Interaction with time	0.002(-0.001,0.005)	1.6	
MUAC(cm)	0.08(-0.01,0.18)		0.12(0.03,0.20)
Interaction with time	0.004(-0.01,0.01)	1	
Polygamous marriage	-0.10(-0.63,0.42)		-0.23(-0.74,0.29)
Interaction with time	0.003(-0.05,0.05)	0	
Autonomy scale (0-5)	-0.10(-0.21,0.01)		-0.15(-0.27,-0.04)
Interaction with time	0.01(0.004,0.02)	7.4	0.01(0.00,0.02)
Use khat and/or alcohol	-0.05(-0.68,0.59)		0.14(-0.39,0.68)
Interaction with time	0.03(-0.03,0.09)	1.2	
obstetric complication	-0.14(-0.52,0.24)		-0.05(-0.41,0.31)
Interaction with time	0.001(-0.001,0.003)	0.8	
Urban residence	0.74(0.16,1.31)		1.67(0.76,2.58)
Interaction with time	0.09(0.04,0.15)	11.2	0.01(-0.08,0.09)
Under 5 children: 0			
1	-0.05(-0.58,0.48)		-0.28(-0.76,0.20)
≥ 2	-0.32(-0.91,0.26)		-0.50(-1.04,0.03)
Interaction with time		3.2	
1 child	-0.04(-0.09,0.01)		
≥ 2 children	-0.01(-0.07,0.04)		
Age of father in years	0.02(-0.01,0.04)		0.02(-0.02,0.05)
Interaction with time	-0.003(-0.01,-0.001)	6.6	-0.001(-0.004, 0.003)
Poverty index (0 - 11)	-0.05(-0.15,0.06)		0.02(-0.14,0.17)
Interaction with time	-0.01(-0.02,-0.01)	9.2	0.01(-0.01,0.02)
Poor sanitary condition (0 -3)	0.11(-0.10,0.32)		0.35(0.09,0.61)
Interaction with time	-0.07(-0.09,-0.05)	43.4	-0.07(-0.09,-0.05)
Level of social support (0 - 4)	0.04(-0.15,0.23)		0.02(-0.14,0.18)
Interaction with time	-0.01(-0.03,0.01)	1.6	
Female gender	-0.81(-1.21,-0.41)		-0.81(-1.15,-0.47)
Interaction with time	-0.005(-0.04,0.03)	0.0	
Not immunised at two months	0.21(-0.20,0.62)		0.34(-0.01,0.69)
Interaction with time	0.01(-0.03,0.05)	0.2	
Severe illness in the first 2 months	-0.08(-0.58,0.41)		-0.03(-0.46,0.39)
Interaction with time	0.01(-0.03,0.06)	0.4	
Birth weight: >=2500g			
<2500g	-1.29(-2.26,-0.32)		-1.75(-2.57,-0.92)
Not measured	-0.03(-0.46,0.41)		-0.49(-0.86,-0.11)
Interaction with time		7.0	
<2500g	-0.06(-0.15,0.03)		
Not measured	-0.05(-0.09,-0.01)		
Non-exclusive breast-feeding at 2 months	-0.28(-0.83,0.27)		-0.13(-0.59,0.34)
Interaction with time	0.04(-0.01,0.09)	2.0	
No pre-lacteal food	-0.10(-1.07,0.88)		-0.66(-1.46,0.15)
Interaction with time	-0.05(-0.14,0.04)	1.2	
Colostrums not given	0.52(0.01,1.04)		0.57(0.12,1.02)
Interaction with time	-0.02(-0.07,0.03)	0.8	
Delayed breast feeding	-0.36(-0.80,0.07)		-0.001(-0.37,0.37)
Interaction with time	0.03(-0.01,0.07)	1.8	

7.4.2 Weight of infants

Summary of results from unadjusted and fully adjusted multilevel growth models of infant weight are presented in table 7.12. Model fit of unadjusted conditional multilevel growth model was significantly improved for selected variables when they were included as predictors of instantaneous weight gain while they were already considered as predictors of attained weight at two months of age: maternal age (change in deviance = 16.8, df = 1), age of the father (change in deviance = 7.4; df = 1), place of residence (change in deviance = 7.2; df = 1), number of previous under five children the family has (change in deviance = 9.4, df =1), poverty index (change in deviance = 10.4; df = 1), and poor sanitary condition scale (change in deviance = 13.8, df = 1)

Before adjusting for potential predictors of an infant weight trajectories an increased maternal height ($\hat{\beta}$ = 0.01; 95%CI: 0.0004, 0.02) or maternal MUAC ($\hat{\beta}$ = 0.06; 95%CI: 0.04, 0.09), having at least one obstetric complication ($\hat{\beta}$ = 0.15; 95%CI: 0.04, 0.26), and living in urban area ($\hat{\beta}$ = 0.28; 95%CI: 0.14, 0.43) were significantly associated with increased attained weight at the age of two months while scoring higher on poverty index ($\hat{\beta}$ = -0.05; 95%CI: -0.07, -0.02), or on poor sanitary condition scale ($\hat{\beta}$ =-0.13; 95%CI; -0.18, -0.08), being a female infant ($\hat{\beta}$ =-0.27; 95%CI: -0.37, -0.17), being low birth weight infant ($\hat{\beta}$ =-0.69; 95%CI: -0.93, -0.15), and delayed initiation of breastfeeding for more than an hour ($\hat{\beta}$ =-0.21; 95%CI: -0.32, -0.10) were inversely related with the attained infant weight at the age of two months. Likewise, having an older mother ($\hat{\beta}$ = -0.002; 95%CI: -0.003, -0.001) or older father ($\hat{\beta}$ =-0.001; 95%CI: -0.002, -0.0003), having one ($\hat{\beta}$ =-0.02; 95%CI: -0.04, -0.01) or more than one sibling ($\hat{\beta}$ =-0.02; 95%CI: -0.03, -0.001) aged less than five years within family compared to families who do not have previous under five child, and scoring higher on poverty index ($\hat{\beta}$ =-0.005; 95%CI: -0.01, -0.002) or on poor sanitary condition scale ($\hat{\beta}$ = -0.01; 95%CI: -0.02, -0.01) were inversely related with the instantaneous rate of weight gain. Infants whose parents reside in urban areas have significantly better instantaneous weight gain ($\hat{\beta}$ = 0.02; 95%CI: 0.01, 0.04) than rurally.

In a fully adjusted model six variables were significantly associated with infant weight: (a) at the age of two months infants whose mothers experienced at least one obstetric complication during current delivery are 150gm heavier than their counterpart ($\hat{\beta}$ =0.15; 95%CI: 0.05, 0.25), (b) an increase in maternal MUAC is significantly associated with increased initial infant weight ($\hat{\beta}$ = 0.06; 95%CI: 0.04, 0.09), (c) increased value on poor sanitary score ($\hat{\beta}$ =-0.07; 95%CI: -0.14, -0.01), being a female infant ($\hat{\beta}$ =-0.31; 95%CI: -0.41, -0.22), to be a low birth weight infant ($\hat{\beta}$ = -0.66; 95%CI: -0.89, -0.43), delayed initiation of breast feeding for more than an hour ($\hat{\beta}$ = -0.15; -0.25, -0.04) were all inversely associated with the weight at the age of two months, and (d) having an older mother ($\hat{\beta}$ = -0.002; 95%CI: -0.003, -0.001) or scoring higher on poor sanitary condition scale ($\hat{\beta}$ = -0.01; 95%CI: -0.02, -0.001) were inversely associated with instantaneous weight gain.

Table 7.12: Interaction of different correlated of infant Weight at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ (95% CI)	Change in deviance	Adjusted $\hat{\beta}$ (95% CI)
Age (years)	-0.004(-0.01,0.005)		-0.005(-0.02, 0.01)
Interaction with time	-0.002(-0.003,-0.001)	16.8	-0.002(-0.003,-0.001)
Height (metres)	0.01(0.0004,0.02)		0.01(-0.001, 0.01)
Interaction with time	-0.00(-0.001,0.001)	0.0	
MUAC(cm)	0.06(0.04,0.09)		0.06(0.04, 0.09)
Interaction with time	0.002(-0.001,0.004)	1.8	
Polygamous marriage	0.04(-0.09,0.18)		0.07(-0.08, 0.21)
Interaction with time	-0.01(-0.02,0.005)	1.8	
Autonomy scale (0-5)	0.01(-0.02,0.04)		-0.01(-0.04, 0.02)
Interaction with time	0.00(-0.003,0.004)	0.2	
Use khat and/or alcohol	0.12(-0.04,0.28)		0.13(-0.02, 0.28)
Interaction with time	0.002(-0.02,0.02)	0.0	
obstetric complication	0.15(0.04,0.26)		0.15(0.05, 0.25)
Interaction with time	0.006(-0.01,0.02)	1.0	
Urban residence	0.28(0.14,0.43)		0.16(-0.07, 0.48)
Interaction with time	0.02(0.01,0.04)	7.2	-0.002(-0.03, 0.02)
Under 5 children: 0			
1	0.06(-0.07,0.20)		0.06(-0.09, 0.20)
≥ 2	0.01(-0.14,0.16)		-0.06(-0.23, 0.10)
Interaction with time		9.4	
1 child	-0.02(-0.04,-0.01)		-0.01(-0.02, 0.01)
≥ 2 children	-0.02(-0.03,-0.001)		-0.004(-0.02, 0.01)
Age of father in years	-0.003(-0.01,-0.0003)		-0.003(-0.01, 0.01)
Interaction with time	-0.001(-0.002,-0.0003)	7.4	0.00(-0.001, 0.001)
Poverty index (0 - 11)	-0.05(-0.07,-0.02)		0.002(-0.04, 0.04)
Interaction with time	-0.005(-0.01,-0.002)	10.4	-0.002(-0.01, 0.002)
Poor sanitary condition (0 -3)	-0.13(-0.18,-0.08)		-0.07(-0.14, -0.01)
Interaction with time	-0.01(-0.02,-0.01)	13.8	-0.01(-0.02, -0.001)
Level of social support (0 - 4)	-0.01(-0.06,0.04)		0.002(-0.04, 0.05,)
Interaction with time	-0.003(-0.01,0.003)	1.0	
Female gender	-0.27(-0.37,-0.17)		-0.31(-0.41, -0.22)
Interaction with time	-0.01(-0.02,0.004)	1.8	
Not immunised at two months	-0.01(-0.11,0.10)		0.02(-0.08, 0.11)
Interaction with time	-0.00(-0.01,0.01)	0.0	
Severe illness in the first 2 months	-0.05(-0.18,0.07)		-0.05(-0.17, 0.07)
Interaction with time	0.01(-0.01,0.02)	0.8	
Birth weight: >=2500g			
<2500g	-0.69(-0.93,-0.15)		-0.66(-0.89, -0.43)
Not measured	0.05(-0.06,0.16)		0.003(-0.10, 0.11)
Interaction with time		1.2	
<2500g	-0.01(-0.04,0.02)		
Not measured	-0.01(-0.02,0.01)		
Non-exclusive breast-feeding at 2 months	0.05(-0.09,0.19)		0.03(-0.10, 0.16)
Interaction with time	-0.001(-0.02,0.01)	0.0	
No pre-lacteal food	-0.02(-0.27,0.23)		-0.05(-0.27, 0.18)
Interaction with time	0.02(-0.03,0.03)	0.0	
Colostrums not given	-0.06(-0.20,0.07)		0.01(-0.11, 0.14)
Interaction with time	-0.01(-0.03,0.00)	3.8	
Delayed breast feeding	-0.21(-0.32,-0.10)		-0.15(-0.25, -0.04)
Interaction with time	-0.001(-0.01,0.01)	0.0	

7.4.3 Length-for-age z score

Information about the predictors of initial length-for-age z and the instantaneous rate of change is summarized in table 7.13. Model fit of unadjusted multilevel growth model was significantly improved (i.e. significant reduction of deviance for number of additional parameters estimated) when maternal age, maternal autonomy score, residence, father's age, poverty index, or poor sanitary condition scale were included as predictors of length-for-age z at two months of age and instantaneous rate of change compared to the model that included them as just predictors of initial length-for-age z.

In unadjusted model infant length-for-age z was significantly better in urban area, if the infant is female or if the colostrums were discarded with the excess of length-for-age z amounting to 0.31 standard deviation units if the family resides in urban area ($\hat{\beta}$ =0.31; 95%CI: 0.07, 0.56), 0.44 standard deviation units if the infant is female ($\hat{\beta}$ =0.44; 95%CI: 0.27, 0.61) and 0.29 standard deviation units if colostrums were discarded ($\hat{\beta}$ = 0.29; 95%CI: 0.06, 0.51). When instantaneous rate of change of length-for-age z is considered as an outcome (a) an increase in maternal age ($\hat{\beta}$ = -0.002; 95%CI: -0.003, -0.001) or father's age ($\hat{\beta}$ =-0.001; 95%CI: -0.002, -0.0003) and scoring higher on poverty index ($\hat{\beta}$ =-0.01; 95%CI: -0.01, -0.001) or scoring higher on poor sanitary condition scale ($\hat{\beta}$ = -0.02; 95%CI: -0.03, -0.02) were significantly and inversely related with instantaneous rate of length-for-age z, and (b) a higher score on maternal autonomy scale ($\hat{\beta}$ =0.004; 95%CI: 0.000, 0.01) and urban residence ($\hat{\beta}$ =0.03; 95%CI: 0.01, 0.05) were significantly associated with higher instantaneous rate of change of length-for-age z score although the statistical significance of maternal autonomy scale was marginal.

In the fully adjusted multilevel growth model (a) increased maternal MUAC ($\hat{\beta}$ = 0.05; 95%CI: 0.01, 0.08), urban residence ($\hat{\beta}$ =0.78; 95%CI: 0.40, 1.17), higher score on a poor sanitary condition scale ($\hat{\beta}$ =0.12; 95%CI: 0.01, 0.23), being a female infant ($\hat{\beta}$ =0.47; 95%CI: 0.33, 0.61), and discarding colostrums ($\hat{\beta}$ =0.25; 95%CI: 0.07, 0.43) were significantly associated with an increased score of a two month old average infant's length-for-age z , (b) being low birtheight infant (($\hat{\beta}$ = -0.57; 95%CI: -0.90, -0.23),

being given pre-lacteal food ($\hat{\beta}$ = -0.38; 95%CI: -0.70, -0.05) and higher score on maternal autonomy scale ($\hat{\beta}$ =-0.06; 95% CI: -0.11, -0.01) were significantly and inversely related with length-for-age z score of a two month old average infant, and (c) increased score on poor sanitary condition scale ($\hat{\beta}$ = -0.03; 95%CI: -0.03, -0.02) was inversely related with the rate of instantaneous change of length-for-age z score.

Table 7.13: Interaction of different correlates of infant length-for-age z at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ *(95% CI)*	*Chance in deviance*	*adjusted* $\hat{\beta}$ *(95% CI)*
Age (years)	0.002(-0.01, 0.02)		-0.01(-0.03,0.01)
Interaction with time	-0.002(-0.003, -0.001)	7.8	-0.001(-0.002,0.001)
Height (metres)	-0.001(-0.02, 0.01)		0.002(-0.01,0.01)
Interaction with time	0.001(-0.001, 0.002)	1.2	
MUAC(cm)	0.03(-0.01, 0.07)		0.05(0.01, 0.08)
Interaction with time	0.002(-0.001, 0.005)	1.4	
Polygamous marriage	0.09(-0.14,0.32)		-0.02(-0.23, 0.19)
Interaction with time	-0.01(-0.02, 0.01)	0.4	
Autonomy scale (0-5)	-0.04(-0.09, 0.01)		-0.06(-0.11, -0.01)
Interaction with time	0.004(0.000, 0.008)	4.6	
Use khat and/or alcohol	-0.04(-0.31, 0.23)		0.07(-0.14, 0.29)
Interaction with time	0.02(-0.004, 0.04)	2.4	
Obstetric complication	-0.10(-0.29, 0.08)		-0.03(-0.18, 0.11)
Interaction with time	0.01(-0.003, 0.03)	2.4	
Urban residence	0.31(0.07, 0.56)		0.78(0.40, 1.17)
Interaction with time	0.03(0.01, 0.05)	7.8	-0.01(-0.04,0.02)
Under 5 children: 0			
1	-0.09(-0.32, 0.13)		-0.08(-0.27, 0.12)
≥ 2	-0.08(-0.33, 0.18)		-0.17(-0.38, 0.05)
Interaction with time		1.8	
1 child	-0.01(-0.03, 0.01)		
≥ 2 children	-0.01(-0.03, 0.01)		
Age of father in years	0.01(-0.003, 0.02)		0.01(-0.01, 0.03)
Interaction with time	-0.001(-0.002, -0.0003)	7.8	-0.001(-0.002, 0.00)
Poverty index (0 - 11)	-0.01(-0.05, 0.03)		0.03(-0.04, 0.09)
Interaction with time	-0.01(-0.01, -0.001)	8.4	0.001(-0.005, 0.01)
Poor sanitary condition (0 -3)	0.03(-0.06, 0.12)		0.12(0.01, 0.23)
Interaction with time	-0.02(-0.03, -0.02)	40.2	-0.03(-0.03, -0.02)
Level of social support (0 - 4)	-0.02(-0.11, 0.07)		-0.02(-0.09, 0.04)
Interaction with time	-0.001(-0.01, 0.01)	0.2	
Female gender	0.44(0.27, 0.61)		0.47(0.33, 0.61)
Interaction with time	0.003(-0.01, 0.02)	0.2	
Not immunised at two months	0.06(-0.11, 0.24)		0.13(-0.01, 0.27)
Interaction with time	0.004(-0.01, 0.02)	0.2	
Severe illness in the first 2 months	-0.13(-0.34, 0.09)		-0.01(-0.18, 0.16)
Interaction with time	0.01(-0.01, 0.02)	0.8	
Birth weight: >=2500g			
<2500g	-0.36(-0.78, 0.06)		-0.57(-0.90, -0.23)
Not measured	-0.09(-0.28, 0.09)		-0.23(-0.39, -0.08)
Interaction with time		5.4	
<2500g	-0.02(-0.06, 0.01)		
Not measured	-0.02(-0.03, -0.001)		
Non-exclusive breast-feeding at 2 months	-0.13(-0.37, 0.11)		-0.01(-0.20, 0.18)
Interaction with time	0.01(-0.01, 0.03)	1.2	
Had pre-lacteal food	-0.20(-0.62, 0.22)		-0.38(-0.70, -0.05)
Interaction with time	-0.01(-0.05, 0.02)	0.4	
Colostrums not given	0.29(0.06, 0.51)		0.25(0.07, 0.43)
Interaction with time	-0.01(-0.03, 0.004)	2.2	
Delayed breast feeding	-0.11(-0.30, 0.08)		0.05(-0.10, 0.20)
Interaction with time	0.01(-0.01, 0.02)	1.0	

7.4.4 Weight-for-age z score

The effects of potential predictors of weight-for-age z score both from unadjusted and adjusted conditional MGMs are summarized in table 7.14. The overall fit of unadjusted conditional MGMs were significantly improved by including parental or household characteristics in the model, namely, age of mother, age of father, residence, number of previous under five siblings in the family, poverty index and poor sanitary condition scale as predictors of instantaneous rate of weight-for-age z change. In other words there were significant reductions in deviance for extra number of new parameters estimated as compared to the models which only allowed the same covariate to predict weight-for-age z score at the two months of age.

In unadjusted conditional MGM (a) infants whose mothers have increased MUAC ($\hat{\beta}$=0.08; 95%CI: 0.05, 0.11) or whose mother had at least one obstetric complication ($\hat{\beta}$=0.21; 95%CI: 0.07, 0.36) or living in urban residence ($\hat{\beta}$=0.34; 95%CI: 0.14, 0.53) or being a female infant ($\hat{\beta}$= 0.35; 95%CI: 0.21, 0.48) have significantly higher weight-for-age z score at the age of two month compared to their counter parts, (b) an increased score on poverty index ($\hat{\beta}$= -0.06; 95%CI: -0.09, -0.02) or on poor sanitary condition scale ($\hat{\beta}$= -0.19; 95%CI: -0.26, -0.12) or being born with low birth weight ($\hat{\beta}$= -1.07; 95%CI: -1.39, -0.74) or delayed initiation of breast feeding for more than an hour ($\hat{\beta}$ = -0.27; 95%CI: -0.42, -0.12) were inversely related with initial weight-for-age z score and, (c) increased maternal age ($\hat{\beta}$= -0.002; 95%CI: -0.01, -0.001) or father's age ($\hat{\beta}$= -0.001; 95%CI: -0.001, -0.0002), having one ($\hat{\beta}$= -0.02; 95%CI: -0.04, -0.01) or more ($\hat{\beta}$= -0.02; 95%CI: -0.03, -0.001) under five siblings, higher score on poverty index ($\hat{\beta}$= -0.004; 95%CI: -0.01, -0.001) or higher score on poor sanitary condition scale($\hat{\beta}$=-0.01; 95%CI: -0.01, -0.001) were inversely and significantly associated with the instantaneous rate of change in weight-for-age z score.

In the fully adjusted model increased maternal age ($\hat{\beta}$=-0.001; 95%CI: -0.003, -0.0001) was the only significant predictor of instantaneous rate and it was inversely associated with the outcome. Urban residence and poverty index did not retain statistical

significance in predicting weight-for-age z score at the age of two months. Increased maternal MUAC ($\hat{\beta}$ =0.08; 95%CI: 0.05, 0.10), being born from mother with at least one obstetric complication ($\hat{\beta}$ =0.17; 95%CI: 0.05, 0.30), and being a female infant ($\hat{\beta}$ = 0.31; 95%CI: 0.19, 0.43) were significantly associated with increased initial weight-for-age z score. However, higher score on poor sanitary condition scale ($\hat{\beta}$ = -0.13; 95%CI: -0.21, -0.05), low birth weight ($\hat{\beta}$ = -0.86, 95%CI: -1.15, -0.58), and delayed initiation of beast feeding for more than an hour ($\hat{\beta}$ = -0.14, 95%CI: -0.27, -0.02) were inversely and significantly related with initial value of weight-for-age z score.

Table 7.14: Interaction of different correlates of infant weight-for-age z at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ *(95% CI)*	*Change in deviance*	*adjusted* $\hat{\beta}$ *(95% CI)*
Age (years)	-0.01(-0.02, 0.003)		-0.01(-0.03, 0.01)
Interaction with time	-0.002(-0.01, -0.001)	13.0	-0.001(-0.003, -0.0001)
Height (metres)	0.01(-0.00, 0.02)		0.01(-0.002, 0.02)
Interaction with time	-0.001(-0.001, 0.003)	1.2	
MUAC(cm)	0.08(0.05, 0.11)		0.08(0.05, 0.10)
Interaction with time	0.00(-0.002, 0.003	0.0	
Polygamous marriage	0.07(-0.11, 0.25)		0.05(-0.13, 0.23)
Interaction with time	-0.01(-0.02, 0.004)	2.0	
Autonomy scale (0-5)	0.01(-0.03, 0.04)		-0.01(-0.04, 0.02)
Interaction with time	0.001(-0.002, 0.004)	0.2	
Use khat and/or alcohol	0.13(-0.09, 0.35)		0.14(-0.05, 0.32)
Interaction with time	0.004(-0.01, 0.02)	0.2	
obstetric complication	0.21(0.07, 0.36)		0.17(0.05, 0.30)
Interaction with time	0.001(-0.01, 0.01)	0.0	
Urban residence	0.34(0.14, 0.53)		0.16(-0.09, 0.42)
Interaction with time	0.02(-0.00, 0.03)	3.8	
Under 5 children: 0			
1	-0.01(-0.19, 0.17)		0.03(-0.16, 0.23)
≥ 2	0.03(-0.17, 0.23)		-0.09(-0.31, 0.12)
Interaction with time		8.4	
1 child	-0.02(-0.04, -0.01)		-0.01(-0.02, 0.01)
≥ 2 children	-0.02(-0.03, -0.001)		-0.002(-0.02, 0.02)
Age of father in years	-0.01(-0.01, 0.003)		-0.002(-0.01, 0.01)
Interaction with time	-0.001(-0.001, -0.0002)	6.2	0.00(-0.001, 0.001)
Poverty index (0 – 11)	-0.06(-0.09, -0.02)		0.01(-0.04, 0.05)
Interaction with time	-0.004(-0.01, -0.001)	6.4	-0.002(-0.01, 0.001)
Poor sanitary condition (0 -3)	-0.19(-0.26, -0.12)		-0.13(-0.21, -0.05)
Interaction with time	-0.01(-0.01, -0.001)	5.2	-0.004(-0.01, 0.003)
Level of social support (0 - 4)	-0.02(-0.08, 0.05)		-0.01(-0.06, 0.05)
Interaction with time	-0.001(-0.007, 0.004)	0.4	
Female gender	0.35(0.21, 0.48)		0.31(0.19, 0.43)
Interaction with time	-0.001(-0.01, 0.01)	0.0	
Not immunised at two months	-0.01(-0.15, 0.13)		0.04(-0.08, 0.17)
Interaction with time	0.001(-0.01, 0.01)	0.2	
Severe illness in the first 2 months	-0.16(-0.33, 0.01)		-0.07(-0.21, 0.08)
Interaction with time	0.01(-0.01, 0.02)	1.2	
Birth weight: >=2500g			
<2500g	-1.07(-1.39, -0.74)		-0.86(-1.15, -0.58)
Not measured	-0.01(-0.15, 0.14)		-0.04(-0.17, 0.09)
Interaction with time		2.6	
<2500g	0.02(-0.01, 0.04)		
Not measured	-0.004(-0.02, 0.01)		
Non-exclusive breast-feeding at 2 months	-0.01(-0.20, 0.18)		0.03(-0.13, 0.19)
Interaction with time	0.004(-0.01, 0.02)	0.4	
No pre-lacteal food	-0.02(-0.36, 0.32)		-0.09(-0.36, 0.19)
Interaction with time	-0.002(-0.03, 0.02)	0.0	
Colostrums not given	-0.09(-0.27, 0.09		-0.002(-0.16, 0.15)
Interaction with time	-0.01(-0.03, 0.001)	3.4	
Delayed breast feeding	-0.27(-0.42, -0.12)		-0.14(-0.27, -0.02)
Interaction with time	0.01(-0.01, 0.02)	1.0	

7.4.5 Stunting

Summary results presented in table 7.15 are on a logit scale and they are obtained from unadjusted and adjusted MGMs where the outcome variable is logit of stunting of infants. Change in deviance is also presented as the comparison of two nested unadjusted conditional MGMs where both of the models have the same predictor of the outcome at two months of age but only one of the two models allow that particular covariate to predict rate of change of the logit of stunting . Only five of the covariates, namely, maternal age, residence, poverty index, poor sanitary condition scale, and birth weight improved the model fit (i.e. resulted in a significant redaction of deviance per extra parameter estimated) and they were individually significant predictors of rate of change of the outcome over time.

Before taking account of other potential risk factors infant gender is the only significant predictor of the logit of stunting at the age of two months. Females have lower risk of being stunted compared to male infants ($\hat{\beta}$ = -0.76; 95% CI: -1.06, -0.45; OR = 0.47, 95%CI: 0.35, 0.64). In other words, at the age of two months the odds of an average male infant to be stunted is 2.14 times that of an average female infant. A one year increased in maternal age ($\hat{\beta}$ = 0.003; 95%CI: 0.001, 0.01; OR = 1.00; 95%CI: 1.00, 1.01;), a one point increase on a poverty index score ($\hat{\beta}$ = 0.01; 95%CI: 0.002, 0.02; OR = 1.002; 95%CI: 1.01, 1.02;), and a one point increase on a poor sanitary condition scale ($\hat{\beta}$ = 0.05 ; 95% ; 0.03, 0.07; OR = 1.05; 95%CI: 1.03, 1.07) were marginally significant predictors of the rate at which logit of stunting changes overtime. Low birth weight ($\hat{\beta}$ = 0.11; 95%CI: 0.01, 0.20; OR = 1.12; 95%CI: 1.01, 1.22) was significantly associated with an increased rate of change in logit of stunting but urban residence (with $\hat{\beta}$ = -0.10, 95%CI: -0.15, -0.04; OR = 0.90; 95%CI: 0.81, 0.96) was significantly and inversely associated with the rate at which logit of stunting changes.

In a fully adjusted conditional MGM significant predictors of initial values of the logit of stunting were an increase of maternal MUAC ($\hat{\beta}$ =-0.08; 95%CI: -0.14, -0.02; OR = 0.92; 95%CI: 0.87, 0.98), urban residence ($\hat{\beta}$ = -0.98; 95%CI: -1.62, -0.23; OR = 0.38;

95%CI: 0.20, 0.79), higher score on poor sanitary condition scale ($\hat{\beta}$ = -0.24; 95%CI: -0.44, -0.03; OR = 0.79; 95%CI: 0.64, 0.97), female gender of an infant ($\hat{\beta}$ = -0.90; 95%CI: -1.15, -0.65; OR = 0.41; 95%CI: 0.32, 0.52), not being vaccinated before two months of age ($\hat{\beta}$ = -0.26; 95% CI: -0.51, -0.01; OR = 0.77; 95%CI: 0.60, 0.99) and being born with low birthweight ($\hat{\beta}$ = 0.69; 95% CI: 0.10, 1.27; OR = 1.99; 95%CI: 1.11, 3.56). The protective effect of better maternal nutritional status, rural residence and increased birthweight are in agreement with the expectation while that of higher score on poor sanitary condition and not being vaccinated before the age of two months are counter intuitive. A one year increase of maternal age ($\hat{\beta}$ = 0.003; 95%CI: 0.0003, 0.01; OR = 1.00; 95%CI: 1.00, 1.01) and scoring one point higher on poor sanitary condition scale ($\hat{\beta}$ =0.04; 95%CI: 0.02, 0.06; OR = 1.04; 95%CI: 1.02, 1.06) were also significant predictors of linear growth of the logit of stunting although the statistical significance of the former was marginal.

Table 7.15: Interaction of different correlates of infant stunting at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ *(95% CI)*	*Change in deviance*	*adjusted* $\hat{\beta}$ *(95% CI)*
Age (years)	0.01(-0.03,0.02)		-0.001(-0.04,0.03)
Interaction with time	0.003(001,0.01)	5.2	0.003(0.0003,0.01)
Height (metres)	0.01(-0.02,0.03)		-0.01(-0.03,0.01)
Interaction with time	-0.002(-0.01,0.001)	2.2	
MUAC(cm)	-0.06(-0.13, 0.01)		-0.08(-0.14,-0.02)
Interaction with time	-0.002(-0.01,0.01)	0.2	
Polygamous marriage	-0.16(-0.57,0.25)		0.08(-0.28,0.45)
Interaction with time	0.02(-0.03,0.07)	0.6	
Autonomy scale (0-5)	0.05(-0.03,0.13)		0.04(-0.03,0.11)
Interaction with time	-0.01(-0.02,0.002)	2.4	
Use khat and/or alcohol	0.03(-0.45,0.51)		-0.19(-0.58,0.20)
Interaction with time	-0.05(-0.10,0.01)	2.8	
obstetric complication	-0.23(-0.55,0.09)		-0.05(-0.31,0.21)
Interaction with time	0.02(-0.02,0.06)	1.2	
Urban residence	-0.18(-0.63,0.28)		-0.98(-1.74,-0.23)
Interaction with time	-0.10(-0.15,-0.04)	12.2	-0.03(-0.12,0.05)
Under 5 children: 0			
1	0.15(-0.25,0.55)		0.17(-0.17,0.52)
≥ 2	-0.002(-0.45,0.45)		0.14(-0.25,0.52)
Interaction with time		1.4	
1 child	0.03(-0.02,0.08)		
≥ 2 children	0.02(-0.03,0.07)		
Age of father in years	-0.005(-0.02, 0.01)		-0.005(-0.03,0.02)
Interaction with time	0.002(-0.001,0.004)	2.2	
Poverty index (0 – 11)	0.01(-0.07,0.09)		-0.0003(-0.12,0.12)
Interaction with time	0.01(0.002,0.02)	5.4	-0.005(-0.02,0.01)
Poor sanitary condition (0 -3)	-0.10(-0.26,0.07)		-0.24(-0.44,-0.03)
Interaction with time	0.05(0.03,0.07)	22.6	0.04(0.02,0.06)
Level of social support (0 – 4)	-0.05(-0.20,0.09)		-0.002(-0.12,0.11)
Interaction with time	0.01(-0.004,0.03)	2.0	
Female gender	-0.76(-1.06,-0.45)		-0.90(-1.15,-0.65)
Interaction with time	-0.02(-0.06,0.02)	1.2	
Not immunised at two months	-0.17(-0.49,0.14)		-0.26(-0.51,-0.01)
Interaction with time	-0.01(-0.05,0.03)	0.2	
Severe illness in the first 2 months	0.24(-0.13,0.61)		-0.07(-0.37,0.23)
Interaction with time	-0.02(-0.07,0.02)	1.2	
Birth weight: >=2500g			
<2500g	0.08(-0.67,0.83)		0.69(0.10,1.27)
Not measured	0.15(-0.18,0.48)		0.40(0.14,0.67)
Interaction with time		7.8	
<2500g	0.11(0.01,0.20)		0.08(-0.02,0.17)
Not measured	0.04(-0.002,0.08)		0.04(-0.003,0.08)
Non-exclusive breast-feeding at 2 months	-0.02(-0.44,0.39)		-0.001(-0.33,0.33)
Interaction with time	0.01(-0.04,0.06)	0.2	
No pre-lacteal food	0.03(-0.70,0.76)		0.29(-0.27,0.85)
Interaction with time	0.03(-0.06,0.12)	0.6	
Colostrums not given	-0.07(-0.47,0.32)		-0.22(-0.54,0.09)
Interaction with time	-0.01(-0.06,0.03)	0.2	
Delayed breast feeding	0.28(-0.05,0.62)		0.03(-0.24,0.29)
Interaction with time	-0.03(-0.07,0.01)	2.4	

7.4.6 Underweight

Taking logit of underweight of infants as an outcome, summary results from unadjusted and fully adjusted conditional MGMs are presented in table 7.16. The first set of coefficients are from conditional MGMs with only one predictor of initial value and a rate of change, and the second set of coefficients are from a fully adjusted conditional MGM where (a) all the covariates are allowed to predict initial value and (b) some of these covariates are also allowed to predict instantaneous rate of change of the logit of underweight. The latter predictors are based on their statistical significance of improved model fit when two nested unadjusted conditional MGMs with and without the predictor of this rate were compared using change in deviance per extra parameters estimated. Seven predictors were included in the fully adjusted model as predictors of instantaneous rate of change, namely, father's age (change in deviance = 4.2, df = 1), maternal age (change in deviance = 6.8, df = 1), maternal MUAC (change in deviance = 4.6, df = 1), number of under five siblings older than the index infant (change in deviance = 6.4, df = 2), residence (change in deviance = 8.2, df = 1), poverty index (change in deviance = 10.4, df = 1), and poor sanitary condition scale (change in deviance = 11.8, df = 1).

In unadjusted MGM increased maternal age ($\hat{\beta}$ =0.005; 95% CI: 0.001, 0.01; OR = 1.01; 95%CI: 1.01, 1.01), increased maternal MUAC ($\hat{\beta}$ = -0.01; 95%CI: -0.02, -0.001; OR = 0.99; 95%CI: 0.98, 1.00), urban residence ($\hat{\beta}$ = -0.12; 95% CI: -0.20, -0.03; OR = 0.89; 95%CI: 0.82, 0.97), one previous under five sibling compared with no such sibling($\hat{\beta}$ = 0.08, 95%CI: 0.02, 0.14; OR = 1.08, 95%CI: 1.02, 1.15), increased father's age ($\hat{\beta}$ = 0.003; 95%CI: 0.00, 0.01; OR = 1.00, 95%CI: 1.00, 1.01), higher score on poverty index ($\hat{\beta}$ = 0.02; 95%CI: 0.01, 0.03; OR = 1.02; 95%CI: 1.01, 1.03) or higher score on poor sanitary condition scale ($\hat{\beta}$ = 0.04; 95%CI: 0.02, 0.07; OR = 1.04; 95%CI: 1.02, 1.07) were significantly associated with instantaneous rate of change of the logit of underweight. Females compared to male infants ($\hat{\beta}$ = -0.99; 95%CI: -1.40; -0.58; OR = 0.37; 95%CI: 0.25, 0.56) and infants whose mothers had at least one obstetric complication during the delivery of index infant compared to those who did not ($\hat{\beta}$ = -0.32, 95%CI: -0.73, -0.10; OR = 0.89; 95%CI: 0.82, 0.97) were less likely to be underweight at the age of two months. Low birth weight infants ($\hat{\beta}$ = 1.68, 95%CI: 0.88,

2.48; OR = 5.37, 95%CI: 3.41, 11.94) and infants within families scoring higher on poor sanitary condition scale ($\hat{\beta}$ = 0.25; 95%CI: 0.04, 0.47; OR = 1.28, 95%CI: 1.04, 1.60) were significantly associated with higher risk of infants' underweight at the age of two months.

In a fully adjusted model female compared to male infants ($\hat{\beta}$ = -0.82; 95% CI: -1.15, -0.48; OR = 0.44; 95%CI: 0.32, 0.62) are at a significantly lower risk and low birth weight infants compared to normal birth weight infants ($\hat{\beta}$ = 1.75; 95%CI: 1.05, 2.44; OR = 5.75; 95%CI: 2.86, 11.47) are at a significantly elevated risk of underweight at the age of two months. A one centimetre increase in maternal MUAC ($\hat{\beta}$ = -0.096; 95% CI; -0.20, 0.01; OR = 0.91; 95%CI: 0.82, 1.01) has marginally non significant protective effect of infant underweight at the age of two months. Scoring a one point higher on poor sanitary condition scale ($\hat{\beta}$ =0.249; 95%CI: -0.0023, 0.5004; OR = 1.29; 95%CI: 1.00, 1.65) and delayed initiation of breast feeding for more than an hour ($\hat{\beta}$ = 0.360, 95% CI: -0.0066, 0.727; OR = 1.43; 95%CI: 0.99, 2.07) are marginally non-significant risk factors for underweight at the age of two months.

None of the predictors of instantaneous rate of change included in the fully adjusted model retain statistical significance which was observed in unadjusted conditional MGM. In assessing the magnitude of the effect size, however, a one centimetre increase in maternal MUAC ($\hat{\beta}$ = -0.011, 95% CI: -0.0023, 0.001; OR=0.99; 95%CI: 0.98, 1.00) and scoring a one point higher on poor sanitary situation scale ($\hat{\beta}$ = 0.0249; 95% CI: -0.003, 0.0528; OR = 1.03; 95%CI: 1.00, 1.05) were marginally non-significant predictors of the instantaneous rate at which logit of underweight of infants change.

Table 7.16: Interaction of different correlates of infant underweight at initial point (two months of age) with time in Butajira Birth cohort, Ethiopia

	$\hat{\beta}$ *(95% CI)*	*Change in deviance*	*adjusted* $\hat{\beta}$ *(95% CI)*
Age (years)	-0.0004(-0.03,0.03)		0.02(-0.03, 0.07)
Interaction with time	0.005(0.001,0.009)	6.8	0.001(-0.004, 0.01)
Height (metres)	-0.02(-0.06,0.01)		-0.02(-0.05, 0.01)
Interaction with time	0.001(-0.003,0.004)	0.0	
MUAC(cm)	-0.09(-0.19,0.01)		-0.10(-0.20, 0.01)
Interaction with time	-0.01(-0.02,-0.001)	4.6	-0.01(-0.02, 0.001)
Polygamous marriage	-0.12(-0.63,0.40)		-0.04(-0.53, 0.45)
Interaction with time	0.03(-0.03,0.09)	0.8	
Autonomy scale (0-5)	-0.01(-0.12,0.09)		0.01(-0.08, 0.10)
Interaction with time	-0.004(-0.02,0.01)	0.4	
Use khat and/or alcohol	-0.21(-0.86,0.45)		-0.29(-0.83, 0.25)
Interaction with time	-0.04(-0.11,0.04)	0.8	
obstetric complication	-0.32(-0.73,-0.10)		-0.26 (-0.61, 0.08)
Interaction with time	-0.03(-0.07,0.02)	1.0	
Urban residence	-0.28(-0.91,0.35)		-0.42(-1.37, 0.54)
Interaction with time	-0.12(-0.20,-0.03)	8.2	0.01(-0.10, 0.12)
Under 5 children: 0			
1	-0.29(-0.80,0.23)		-0.14(-0.73,0.45)
≥ 2	-0.20(-0.78,0.37)		0.21(-0.44,0.85)
Interaction with time		6.4	
1 child	0.08(0.02,0.14)		0.03(-0.04,0.09)
≥ 2 children	0.04(-0.02,0.11)		0.01(-0.06,0.08)
Age of father in years	0.002(-0.02,0.03)		0.00(-0.04,0.04)
Interaction with time	0.003(0.00,0.01)	4.2	0.001(-0.003,0.005)
Poverty index (0 – 11)	0.02(-0.09,0.12)		-0.101(-0.25,0.05)
Interaction with time	0.02(0.01,0.03)	10.4	0.01(-0.01,0.03)
Poor sanitary condition (0 -3)	0.25(0.04,0.47)		0.25(-0.002,0.50)
Interaction with time	0.04(0.02,0.07)	11.8	0.02(-0.003,0.05)
Level of social support (0 - 4)	-0.005(-0.20,0.20)		0.02(-0.14,0.18)
Interaction with time	0.01(-0.01,0.03)	1.2	
Female gender	-0.99(-1.40,-0.58)		-0.82(-1.15,-0.48)
Interaction with time	0.02(-0.03,0.07)	0.6	
Not immunised at two months	-0.06(-0.46,0.35)		-0.14(-0.48,0.20)
Interaction with time	-0.01(-0.05,0.04)	0.0	
Severe illness in the first 2 months	0.22(-0.26,0.70)		0.18(-0.22,0.58)
Interaction with time	0.01(-0.04,0.07)	0.2	
Birth weight: >=2500g			
<2500g	1.68(0.88,2.48)		1.75(1.05,2.44)
Not measured	-0.24(-0.68,0.20)		-0.05(-0.42,0.31)
Interaction with time		3.8	
<2500g	0.07(-0.03,0.16)		
Not measured	0.04(-0.01,0.09)		
Non-exclusive breast-feeding at 2 months	0.12(-0.42,0.66)		-0.09(-0.54,0.36)
Interaction with time	-0.02(-0.09,0.04)	0.6	
No pre-lacteal food	-0.10(-1.08,0.88)		0.14(-0.63,0.92)
Interaction with time	0.03(-0.08,0.14)	0.4	
Colostrums not given	0.24(-0.27,0.75)		-0.03(-0.45,0.40)
Interaction with time	0.02(-0.03,0.08)	0.6	
Delayed breast feeding	0.41(-0.04,0.86)		0.36(-0.01,0.73)
Interaction with time	0.02(-0.04,0.07)	0.4	

7.4.7 Summary of findings from unadjusted and fully adjusted MGM

7.4.7.1 Summary findings from unadjusted conditional MGM

Common risk factors to the six growth outcome measures: In models with only one level-2 predictor of initial value and rate of change, females compared to male infants were significantly shorter, lighter, had increased length-for-age z, had increased weight-for-age z , and less likely to be stunted or underweight at the age of two months. Low birthweight infants had significantly shorter length, lighter weight, smaller weight-for-age z, and increased risk of underweight at the age two months but not significantly different from normal birthweight infants in terms of length-for-age z and logit of stunting. Moreover, it was positively associated with the rate at which logit of stunting changes with infant's age. Increased maternal age, higher score on poverty scale or higher score on poor sanitary condition scale were significantly associated with smaller rate with which the six growth outcome variables change. Residential place and age of the father were also significantly associated with the rate of change of all the growth outcomes although the statistical significance of the association of residence with weight-for-age z and the association of father's age with the logit of stunting were marginal.

Risk factors specific to length, length-for-age z and stunting: Denial of colostrum was significantly associated with an increased length and a larger length-for-age z score of a two month old average infant but not with the logit of stunting or with the rate at which the length, length-for-age z or logit of stunting of a two month old average infant changes. Higher score on maternal autonomy scale was significantly associated with an increased rate of change of length or length-for-age z over time but not with the rate that the logit of stunting changes.

Risk factors specific to weight, weight-for-age z and underweight: Having at least one obstetric complication was significantly associated with increased weight, larger values of weight-for-age z and reduced risk of underweight of an average two month old infant. Similarly, larger maternal MUAC and delayed initiation of breast feeding were significantly associated with an increased weight and larger weight-for-age z but not

significantly associated with the logit of underweight of a two month old average infant. Rate of change of weight and weight-for-age z were significantly and inversely associated with having at least one previous sibling below the age of five and the rate of change of the logit of underweight was significantly and positively associated with the presence of one siblings aged less than 5 years. Similarly, increased maternal MUAC was significantly and inversely associated with the rate with which logit of underweight changes.

7.4.7.2 Findings from fully adjusted conditional MGM

Common risk factors to the six growth outcome variables: In a fully adjusted conditional MGM infant gender, birthweight and poor sanitary condition scale were significant predictors of infant growth attainment on all the six domains at two months of age although the statistical significance of the association between poor sanitary condition scale and logit of underweight was marginal. MUAC was significant predictor for the growth attainment of all growth outcomes but not for logit of underweight. Maternal age marginally affected the rate of change of all six growth outcome. Poor sanitary condition scale was significant predictor of the rate of linear change of all growth outcomes with the exception of weight-for-age Z and logit of underweight.

Risk factors specific to length, length-for-age and stunting: At the age of two months place of residence had significant impact on the level of attainment of all three infant growth measures. Increased maternal autonomy score and getting colostrum were significantly and inversely associated with length and length-for-age z attained by a two month old average infant but not with logit of stunting. Moreover, pre-lacteal feeding and immunization before two months of age were significantly associated with the level of length-for-age z and logit of stunting, respectively, at the age of two months. Maternal autonomy was significant predictor of the rate of linear change in length.

Risk factors specific to weight, weight-for-age z and underweight: Infants whose mothers had at least one obstetric complication compared to those who had not, had attained significantly better weight and weight-for-age z, and less likely to be underweight at the age of two months. Delayed initiation of breastfeeding for more than

an hour was significantly associated with lighter weight and relatively lower weight-for-age z, and it was not significantly association with underweight

7.5. Conditional and unconditional MGM – The effect of CMD on initial growth and on its rate of change over time

7.5.1 Introduction

In previous sections of this chapter (a) we decided on the best fitting polynomial function to describe each growth outcome, (b) we investigated adjusted and unadjusted effects of pre-determined risk factors on initial growth of each growth outcome in a model with no predictor of the rate of change, (c) taking one pre-specified risk factor at a time we investigated potential predictors of the rate of change for each growth outcome, and (d) in a model with all covariates included as predictors of initial growth of each growth outcome and selected covariates are also considered as predictors of the rate of change we estimated adjusted effects of different risk factors on initial value and rate of change of each growth outcome.

In this section we investigated the effect of perinatal CMD on infant growth introducing measures of perinatal CMD as one predictor in MGMs used in previous sections of this chapter. As the starting point we used the best fitting polynomial function for each growth outcome determined in the first section of this chapter. The first step was to evaluate unadjusted effects of perinatal CMD on initial value and on the rate of change of each growth outcome by considering one of the perinatal CMD indicator variables (prevalence antenatal, prevalence postnatal, or course of perinatal CMD as a four level variable where no-perinatal CMD was taken as a reference category) as the only predictor of growth parameters. The second step was to investigate adjusted effects of perinatal CMD on initial value of each growth outcome. In the later case two separate models were considered for each growth outcome: (a) conditional MGM with all covariates include as predictors of initial value ignoring predictors of rate of change and (b) the same model as (a) except selected covariates are also considered as predictors of rate of change over time.

Since the effects of other covariates on growth parameters are already summarized in previous sections of this chapter the focus of this section is only on the effects of perinatal CMD. The result will be summarized below as follows

- Unadjusted effects of perinatal CMD on initial value and on the rate of change of each growth outcome
- Adjusted effects of perinatal CMD on initial value of each growth outcome when all covariates are considered as predictors of initial value but none of them are considered as predictors of the rate of change
- Adjusted effects of perinatal CMD on initial value of each growth outcome when all covariates are considered as predictors of initial value and selected covariates are also considered as predictors of the rate of change

7.5.2 **Unadjusted effects of CMD on initial growth and on its rate of change over time**

7.5.2.1 The effect of CMD on infant length

Unadjusted effects of antenatal and postnatal CMD on initial infant length and the rate at which infants gain their length over time are summarized in table 7.17. When antenatal CMD was considered as the only predictor of initial length (table 7.17 column 2) the overall fit of the best fitting unconditional MGM (table 7.1 model 4) was not significantly improved. However, replacing antenatal CMD by postnatal CMD as the only predictor of initial length (table 7.17 column 4) significantly improved the overall model fit of the best fitting unconditional MGM (table 7.1 model 4) (change in deviance = 72.8, df = 1; change in AIC = 70.7). Relaxing model assumption and allowing maternal CMD to predict initial value and rate of change of length over time (table 7.17 column 3 for antenatal CMD and column 5 for postnatal CMD) did not result in a significant improvement of the model fit.

Examining the statistical significance of the fixed effects of CMD across the four models (table 7.17) there was no significant effect of antenatal CMD or postnatal CMD on initial length or on the rate at which infants gain length over time. However, examination of the random components showed that there was significant within individual variability in all the models. Similarly, there was significant variability between infants initial length as well as the rate at which infants gain length. The amount of variability in initial value and in the rate of change explained by antenatal or postnatal CMD was extremely low. This implies that other factors are more important than antenatal or postnatal CMD in explaining why infants gain length differently during infancy.

Table 7.17: Effect of CMD on the trajectory of length of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of length Estimate (95% CI)	*Antenatal CMD as predictor of random intercept and random slope of length Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of length Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of length Estimate (95% CI)*
Fixed Effects (main effects)				
Intercept (status at two months)	57.55(57.3, 57.8)	57.54(57.29, 57.78)	57.57(57.33, 57.81)	57.57(57.33, 57.81)
Age (linear term)	1.61(1.55, 1.66)	1.61(1.55, 1.67)	1.61(1.55, 1.66)	1.60(1.55, 1.66)
Age^2(Quadratic term)	-0.03(-0.04, -0.03)	-0.034(-0.04, -0.03)	-0.03(-0.04, -0.03)	-0.03(-0.04, -0.03)
CMD in pregnancy (SRQ-20 >= 6)	0.05(-0.47, 0.58)	0.18(-0.45, 0.80)		
CMD postnatal (SRQ-20 >= 6)			-0.30(-1.11, 0.51)	-0.42(-1.39, 0.55)
Fixed effect (interaction effect)				
Antenatal CMD by age		-0.02(-0.08, 0.04)		
Postnatal CMD by age				0.02(-0.07, 0.11)
Variance component				
Level 1 :				
Within person	2.89(2.81, 2.97)	2.89(2.81, 2.97)	2.89(2.81, 2.97)	2.89(2.81, 2.97)
Level 2				
Status at two months of age *Linear term*	2.16(1.94, 2.40)	2.16(1.94, 2.40)	2.15(1.94, 2.39)	2.15(1.94, 2.39)
Standard deviation	0.15(0.12, 0.18)	0.15(0.12, 0.18)	0.15(0.12, 0.18)	0.15(0.12, 0.18)
correlation with at two month status	-0.10(-0.31, 0.11)	-0.10(-0.31, 0.11)	-0.11(-0.31, 0.10)	-0.11(-0.31, 0.10)
Goodness of fit				
Log likelihood	-11821.4	-11821.2	-11785.0	-11784.9
AIC	23658.8	23660.3	23586.1	23587.9
BIC	23710.0	23717.9	23637.3	23645.5

7.5.2.2 The effect of CMD on infant weight

Unadjusted effects of antenatal and postnatal CMD on the trajectory of infant weight are summarized in table 7.18. As the only predictor of initial weight antenatal CMD did not improve the overall fit of the best fitting unconditional growth model summarized in table 7.2 model 4. However, replacing antenatal CMD by postnatal CMD significantly improved the overall fit of the model summarized in table 7.2 model 4 (change in deviance = 34.6, df = 1; Change in AIC = 32.5). When each of them were also treated as predictors of rate of change of weight, the overall fit of conditional MGM in which they were the only predictors of initial weight (table 18 column 2 for antenatal CMD and column 4 for postnatal CMD) was not significantly improved

Among the fixed effects across the four models (table 7.18) the only significant effect was that of postnatal CMD on initial weight of infants. In a more restrictive model where postnatal CMD was the predictor of only initial weight, infants whose mothers had postnatal CMD weighed 260gm (95%CI: 30gm , 500 gm) more than those infants whose mothers did not have the same experience. In a less restrictive model which allows postnatal CMD to be also a predictor of rate of change in weight, the difference in initial weight was increased to 280gm (95% CI: 30gm , 520 gm). However, postnatal CMD was not significant predictor of the rate of change of weight over time.

There was significant within individual variability over time and significant between individual variability in initial weight as well as in the rate at which infants gain weight. Postnatal CMD explained 4.2% of the overall variability of the initial weight among infants but antenatal CMD did not reduce any measurable amount of the same variability. CMD at either time points explained extremely low proportion of the variability of rate at which weight of infants change over time.

Table 7.18: Effect of CMD on the trajectory of weight of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of weight Estimate (95% CI)	*Antenatal CMD as predictor of random intercept and random slope of weight Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of weight Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of weight Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	5.18(5.11, 5.24)	5.17(5.11, 5.24)	5.17(5.11, 5.23)	5.17(5.11, 5.23)
Age (linear term)	0.39(0.38, 0.41)	0.39(0.38, 0.41)	0.39(0.38, 0.41)	0.39(0.38, 0.41)
Age^2(Quadratic term)	-0.007(-0.008,-0.006)	-0.007(-0.008,-0.006)	-0.007(-0.008,-0.006)	-0.007(-0.008,-0.006)
CMD in pregnancy(SRQ-20 >= 6)	0.039(-0.113, 0.192)	0.077(-0.083, 0.237)		
CMD postnatal			0.26(0.03, 0.50)	0.28(0.03, 0.52)
Fixed effect (Interaction effect)				
Antenatal CMD by age		-0.013(-0.031,0.004)		
Postnatal CMD by age				-0.005(-0.032,0.022)
Variance component				
Level 1 :				
Within person	0.83(0.81, 0.85)	0.83(0.81, 0.85)	0.83(0.81, 0.85)	0.83(0.81, 0.85)
Level 2				
Status at two months of age	0.47(0.40, 0.54)	0.47(0.40, 0.54)	0.46(0.40, 0.54)	0.46(0.40, 0.54)
Linear term				
Standard deviation	0.05(0.04, 0.06)	0.05(0.04, 0.06)	0.05(0.04, 0.06)	0.05(0.04, 0.06)
correlation with at two month status	0.74(0.18, 0.94)	0.74(0.17, 0.94)	0.74(0.17, 0.94)	0.74(0.17, 0.94)
Goodness of fit				
Log likelihood	-6368.2	-6367.0	-6351.0	-6351.0
AIC	12752.4	12752.1	12718.1	12719.9
BIC	12803.6	12809.7	12769.3	12777.6

7.5.2.3 The effect of CMD on infant length-for-age z

Unadjusted effects of antenatal CMD on initial length-for-age z of infants (column 2) and the rate at which their length-for-age z change over time (column 4), and corresponding effects of postnatal CMD (column 3 & 5) are summarized in table 7.19. Having antenatal CMD as the only predictor of initial length-for-age z did not improve the overall fit of the best fitting unconditional growth model summarized in table 7.3 model 4. However, replacing antenatal CMD by postnatal CMD significantly improved the overall model fit (change in deviance = 53.0, df = 1; Change in AIC = 51.0). When each of them are also treated as predictors of how the rate of length-for-age z changes, the overall fit of conditional MGM in which they were the only predictors of initial length (table 7.19 column 2 for antenatal CMD and column 4 for postnatal CMD) was not significantly improved

Across the four models presented in table 7.19 there was no significant effect of antenatal or postnatal CMD as predictors of initial length-for-age z or the rate at which length-for-age z changes over time. The random component of all the models showed that there was significant within individual variability. Similarly, there was significant between individual variability in initial length-for-age z as well as the rate at which length-for-age z changed. The amount of variability in the initial value or the rate of change in length-for-age z explained by antenatal CMD or postnatal CMD was extremely low. For example, antenatal CMD explained 1.9% of the variability that existed within the initial length-for-age z.

Table 7.19: Effect of CMD on the trajectory of length-for-age of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of length-for-age Estimate (95% CI)	*Antenatal CMD as predictor of random intercept and random slope of length-for-age Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of length-for-age Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of length-for-age Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	-0.23(-0.33, -0.12)	-0.24(-0.34, -0.13)	-0.22(-0.32, -0.12)	-0.22(-0.32, -0.11)
Age (linear term)	-0.26(-0.28, -0.24)	-0.26(-0.28, -0.23)	-0.26(-0.28, -0.24)	-0.26(-0.28, -0.24)
Age^2(Quadratic term)	0.008(0.006, 0.009)	0.008(0.006, 0.009)	0.008(0.006, 0.009)	0.008(0.006, 0.009)
CMD in pregnancy(SRQ-20 >= 6)	-0.05(-0.27, 0.16)	0.03(-0.24, 0.30)		
CMD postnatal			-0.31(-0.64, 0.03)	-0.40(-0.81, 0.02)
Fixed effect (Interaction effect)				
Antenatal CMD by age		-0.01(-0.03, 0.01)		
Postnatal CMD by age				0.01(-0.02, 0.05)
Variance component				
Level 1 :				
Within person	1.16(1.13, 1.20)	1.16(1.13, 1.20)	1.16(1.13, 1.20)	1.16(1.13, 1.20)
Level 2				
Status at two months of age *Linear term*	0.98(0.90, 1.07)	0.98(0.90, 1.07)	0.97(0.89, 1.07)	0.97(0.89, 1.07)
Standard deviation	0.05(0.03, 0.06)	0.05(0.03, 0.06)	0.05(0.03, 0.06)	0.05(0.03, 0.06)
correlation with at two month status	-0.29(-0.47, -0.09)	-0.29(-0.47, -0.09)	-0.29(-0.47, -0.09)	-0.29(-0.47, -0.09)
Goodness of fit				
Log likelihood	-7671.0	-7670.5	-7644.6	-7644.3
AIC	15358.0	15359.0	15305.2	15306.7
BIC	15409.1	15416.5	15356.3	15364.2

7.5.2.4 The effect of CMD on infant weight-for-age z

Unadjusted effects of antenatal CMD on initial weight-for-age z (column 2), on the rate at which their weight-for-age z changes over time (column 3), and corresponding effects of postnatal CMD (columns 4 & 5) are summarized in table 7.20. When antenatal CMD was the only predictor of initial weight-for-age the overall fit of the best fitting unconditional growth model summarized in table 7.4 model 4 was not significantly improved. However, postnatal CMD significantly improved the overall model fit (change in deviance = 35.4.0, df = 1; Change in AIC = 33.5). When each of them were also treated as predictors of rate at which weight-for-age z changes , the overall fit of conditional MGM in which they were the only predictors of initial length (table 7.20 column 2 for antenatal CMD and column 4 for postnatal CMD) was not significantly improved

Across the four models summarized in table 7.20 where maternal CMD was used as predictor of initial value or the rate at which weight-for-age z changes there was no significant effect of CMD on initial value or rate of change. In the random part of all the models there was significant within individual variability. Similarly, there was significant between individual variability in initial weight-for-age z as well as in the rate at which weight-for-age z changed. The amount of variability in initial value and in the rate at which weight-for-age z changes explained by antenatal CMD or postnatal CMD was negligible.

Table 7.20: Effect of CMD on the trajectory of weight-for-age of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of weight-for-age *Estimate (95% CI)*	*Antenatal CMD as predictor of random intercept and random slope of weight-for-age* *Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of weight-for-age* *Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of weight-for-age* *Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	-0.54(-0.62, -0.46)	-0.55(-0.64, -0.47)	-0.55(-0.63, -0.47)	-0.56(-0.64, -0.48)
Age (linear term)	-0.14(-0.15, -0.12)	-0.14(-0.15, -0.12)	-0.14(-0.15, -0.12)	-0.14(-0.15, -0.12)
Age^2(Quadratic term)	0.007(0.006, 0.008)	0.007(0.006, 0.008)	0.007(0.006, 0.008)	0.007(0.006, 0.008)
CMD in pregnancy	-0.04(-0.22, 0.15)	0.07(-0.15, 0.28)		
CMD postnatal			0.18(-0.12, 0.47)	0.27(-0.06, 0.60)
Fixed effect (Interaction effect)				
Antenatal CMD by age		-0.017(-0.034, 0.000)		
Postnatal CMD by age				-0.02(-0.04, 0.01)
Variance component				
Level 1 :				
Within person	0.90(0.88, 0.93)	0.90(0.88, 0.92)	0.90(0.88, 0.93)	0.90(0.88, 0.93)
Level 2				
Status at two months of age	0.81(0.75, 0.88)	0.81(0.75, 0.88)	0.81(0.74, 0.88)	0.81(0.74, 0.88)
Linear term				
Standard deviation	0.04(0.03, 0.05)	0.04(0.03, 0.05)	0.04(0.03, 0.05)	0.04(0.03, 0.05)
correlation with at two month status	0.04(-0.21, 0.28)	0.05(-0.20, 0.29)	0.05(-0.20, 0.29)	0.05(-0.20, 0.29)
Goodness of fit				
Log likelihood	-6742.6	-6740.7	-6725.0	-6724.3
AIC	13501.2	13499.4	13465.9	13466.7
BIC	13552.4	13557.0	13517.1	13524.3

7.5.2.5 The effect of CMD on infant logit of stunting

Unadjusted effects of antenatal CMD on initial logit of stunting (column 2), the rate at which their logit of stunting changes over time (column 3), and corresponding effects of postnatal CMD (columns 4 & 5) are summarized in table 7.21. As the only predictor of initial logit of stunting antenatal CMD did significantly improve (change in deviance = 2.6, df = 1, Change in AIC = 2.8) of the overall fit of the best fitting unconditional growth model summarized in 7.5 model 4. However, replacing antenatal CMD by postnatal CMD significantly improved the overall model fit (change in deviance = 34.6.4.0, df = 1; Change in AIC = 9.0). When each of them were also treated as predictors of rate at which logit of stunting changes, the overall fit of conditional MGM in which they were the only predictors of initial logit of stunting (table 7.21 column 2 for antenatal CMD and column 4 for postnatal CMD) was not significantly improved

As a predictor of initial logit of stunting or the rate at which logit of stunting changes over time, antenatal or postnatal CMD did not show significant effect in each of the four models presented in table 7.21. In all the models there was significant between individual variability in initial logit of stunting as well as the rate at which logit of stunting changed over time. Antenatal or postnatal CMD explained 29.7% of the total variability in the initial value of the logit but neither of them explained any meaningful percent of the variability in the rate at which the logit of stunting changes.

Table 7.21: Effect of CMD on the trajectory of stunting of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of logit of stunting Estimate (95% CI)	*Antenatal CMD as predictor of random intercept and random slope of logit of stunting Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of logit of stunting Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of logit of stunting Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	-2.03(-2.24, -1.83)	-2.02(-2.23, -1.82)	-2.03(-2.23, -1.83)	-2.02(-2.22, -1.81)
Age (linear term)	0.19(0.17, 0.22)	0.19(0.17, 0.22)	0.19(0.17, 0.22)	0.19(0.17, 0.21)
Age^2(Quadratic term)				
CMD in pregnancy	0.05(-0.32, 0.41)	-0.03(-0.51, 0.45)		
CMD postnatal			0.12(-0.44, 0.68)	-0.15(-0.90, 0.60)
Fixed effect (Interaction effect)				
Antenatal CMD by age		0.01(-0.04, 0.07)		
Postnatal CMD by age				
Variance component				0.05(-0.04, 0.14)
Level 1 :				
Within person				
Level 2				
Status at two months of age *Linear term*	0.78(0.47, 1.30)	0.78(0.47, 1.30)	0.78(0.47, 1.30)	0.77(0.46, 1.30)
Standard deviation	0.12(0.09, 0.17)	0.12(0.09, 0.17)	0.12(0.09, 0.17)	0.12(0.09, 0.17)
correlation with at two month status	0.47(-0.53, 0.93)	0.48(-0.54, 0.93)	0.48(-0.54, 0.93)	0.49(-0.56, 0.94)
Goodness of fit				
Log likelihood	-2479.0	-2478.9	-2473.2	-2472.5
AIC	4970.1	4971.8	4958.3	4959.1
BIC	5008.4	5016.6	4996.7	5003.8

7.5.2.6 The effect of CMD on infant logit of underweight

Unadjusted effects of antenatal CMD on initial logit of underweight (column 2), the rate at which logit of underweight changes (column 3), and corresponding effects of postnatal CMD (column 4 & 5) are summarized in table 7.22. As the only predictor of initial logit of underweight antenatal CMD did significantly improve (change in deviance = 13.0 df = 1, Change in AIC = 14.9) the overall fit of the best fitting unconditional growth model summarized in table 7.6 model 4. However, replacing antenatal CMD by postnatal CMD did not improved the overall model fit (change in deviance = 0.6, df = 1; Change in AIC = 1.4). When each of them were also treated as predictors of rate of change of the logit of underweight, the overall fit of conditional MGM in which they were the only predictors of initial logit of underweight (table 7.22 column 2 for antenatal CMD and column 4 for postnatal CMD) was not significantly improved

Across the four models presented in table 7.22 there was no significant effect of antenatal or postnatal CMD as predictor of initial value of the logit of underweight or the rate at which the logit of underweight changes over time. There was significant between individual variability in initial value of the logit of underweight as well as the rate at which it changed over time. Antenatal CMD and postnatal CMD explained 4.0% and 7.9% of the variability in initial value of the logit of underweight, respectively, when they were included as the only predictors. When the model assumption was relaxed and they were also allowed to predict rate of change in the logit of underweight antenatal CMD explained 5.3% and postnatal CMD explained 6.6% of the variability in the initial value of the logit of underweight

Table 7.22: Effect of CMD on the trajectory of underweight of infant data from the P-MaMiE study

	Antenatal CMD as predictor of random intercept of logit of underweight Estimate (95% CI)	*Antenatal CMD as predictor of random intercept and random slope of logit of underweight Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept of logit of underweight Estimate (95% CI)*	*Postnatal CMD as predictor of random intercept and random slope of logit of underweight Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	-2.87(-3.33, -2.41)	-2.85(-3.30, -2.39)	-2.80(-3.25, -2.36)	-2.80(-3.24, -2.35)
Age (linear term)	0.26(0.17, 0.34)	0.25(0.16, 0.34)	0.25(0.16, 0.34)	0.25(0.16, 0.34)
Age^2(Quadratic term)	-0.02(-0.02, -0.01)	-0.02(-0.02, -0.01)	-0.02(-0.02, -0.01)	-0.02(-0.02, -0.01)
CMD in pregnancy	0.16(-0.35, 0.66)	-0.01(-0.62, 0.61)		
CMD postnatal			-0.66(-1.49, 0.17)	-1.05(-0.22, 0.08)
Fixed effect (Interaction effect)				
Antenatal CMD by age		0.03(-0.04, 0.10)		
Postnatal CMD by age				0.07(-0.05, 0.19)
Variance component				
Level 1 :				
Within person				
Level 2				
Status at two months of age	1.46(1.07, 2.00)	1.45(1.06, 1.99)	1.43(1.04, 1.97)	1.44(1.05, 1.97)
Linear term				
Standard deviation	0.17(0.12, 0.23)	0.17(0.12, 0.23)	0.17(0.12, 0.24)	0.17(0.12, 0.23)
correlation with at two month status	0.16(-0.32, 0.58)	0.18(-0.32, 0.60)	0.17(-0.32, 0.59)	0.17(-0.33, 0.59)
Goodness of fit				
Log likelihood	-1913.4	-1913.0	-1906.6	-1906.1
AIC	3840.8	3842.0	3827.3	3828.1
BIC	3885.6	3893.2	3872.1	3879.3

7.5.2.7 The effect of the course of CMD on infant growth

Previously we investigated the effects of antenatal CMD (i.e. prevalence) and postnatal CMD (i.e. prevalence) on each growth outcome. In here we evaluated the effects of the course of perinatal CMD as a four level variable (i.e. no exposure to perinatal CMD, exposure to CMD only during pregnancy, exposure to CMD only in the first two month postnatal , exposure to persistent CMD from pregnancy to two month postnatal period) on infant growth.

The overall model fit of the baseline random coefficient quadratic growth model (random intercept and random instantaneous rate of change models) were improved when the four category CMD variable was considered as the only level-2 predictor of initial value of length (change deviance = 73.1, df = 3, change in AIC = 67), weight (change deviance = 35, df = 1, change in AIC = 29), length-for-age z (change deviance = 53, df = 1, change in AIC = 47), weight-for-age z (change deviance = 31, df = 1, change in AIC = 31), or logit of stunting (change deviance = 9.4 df = 1, change in AIC = 5.3) but did not improve the fit as the predictor of the initial value of the logit of underweight (change deviance = 2.0, df = 1, change in AIC = 3.9).

Relaxing model assumption and allowing course of CMD to also predict rate of change in length (change deviance = 1.2, df = 1, change in AIC = 1.4), weight (change deviance = 2.4, df = 1, change in AIC = 3.5), length-for-age z (change deviance = 2.8, df = 1, change in AIC = 3.2), weight-for-age z (change deviance = 4.0, df = 1, change in AIC = 1.9), logit of stunting (change deviance = 2.4 df = 1, change in AIC = 3.7) or logit of underweight (change deviance = 1.6, df = 1, change in AIC = 4.3) did not improve the overall fit of the models.

Examining the fixed effect parts of all the models (tables 7.23- 7.25) the only significant effect of the course of CMD was related to initial weight of infants whose mothers had experienced chronic CMD. Infants whose mothers had experienced chronic CMD weighed more at the age of two months compared to infants whose mothers had never experienced CMD during perinatal period. There was no significant difference in the rate at which each group of infants gain weight over time between infants whose mothers had experienced perinatal CMD and infants whose mothers had no such experience.

Table 7.23: Effect of course of CMD on the trajectory of height and weight of infant data from the P-MaMiE study

	Course of CMD as predictor of random intercept o f length *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of length* *Estimate (95% CI*	*Course of CMD as predictor of random intercept o f weight* *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of weight* *Estimate (95% CI*
Fixed Effects (Main effects)				
Intercept (status at two months)	57.46(57.31,57.80)	57.55(57.30,57.80)	5.17(5.11, 5.24)	5.17(5.10, 5.23)
Age (linear term)	1.61(1.55, 1.66)	1.61(1.55, 1.66)	0.39(0.38, 0.41)	0.39(0.38, 0.41)
Age^2(Quadratic term)	-0.03(-0.04, -0.03)	-0.03(-0.04, -0.03)	-0.007(0.008, -0.006)	-0.007(0.008, -0.006)
Course of CMD				
Both times negative	Ref	Ref		
Positive only at pregnancy	0.10(-0.48, 0.69)	0.23(-0.47, 0.94)	-0.03(-0.20, 0.14)	-0.01(-0.17, 0.19)
Incident postnatal	-0.49(-1.71, 0.73)	-0.81(-2.28, 0.65)	0.18(-0.18, 0.53)	0.20(-0.18, 0.57)
persistent CMD	-0.14(-1.20, 0.93)	-0.09(-1.36, 1.18)	0.32(0.01, 0.63)	0.34(0.01, 0.66)
Fixed effect (Interaction effect)				
Positive only at pregnancy with time		-0.02(-0.09, 0.04)		-0.02(-0.035, 0.004)
Incident postnatal with time		0.05(-0.08, 0.19)		-0.008(-0.05, 0.03)
persistent CMD with time		-0.008(-0.12, 0.11)		-0.005(-0.04, 0.03)
Variance component				
Level 1 : Within person	2.89(2.81, 2.97)	2.89(2.81, 2.97)	0.83(0.81, 0.85)	0.83(0.81, 0.85)
Level 2 : Status at two months of age	2.15(1.93, 2.39)	2.15(1.93, 2.39)	0.46(0.40, 0.54)	0.46(0.40, 0.54)
Linear term: Standard deviation	0.15(0.12, 0.18)	0.15(0.12, 0.18)	0.05(0.04, 0.06)	0.05(0.04, 0.06)
correlation with at two month status	-0.11(-0.31, 0.10)	-0.11(-0.31, 0.11)	0.74(0.17, 0.94)	0.75(0.16, 0.94)
Goodness of fit : Log likelihood	-11784.9	-11784.3	-6350.8	-6349.6
AIC	23589.8	23594.7	12721.6	12725.1
BIC	23653.7	23677.8	12785.6	12808.4

Table 7.24: Effect of course of CMD on the trajectory of height-for-age and weight-for-age of infant data from the P-MaMiE study

	Course of CMD as predictor of random intercept o f length-for-age *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of length-for-age* *Estimate (95% CI*	*Course of CMD as predictor of random intercept o f weight-for-age* *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of weight-for-age* *Estimate (95% CI*
Fixed Effects (Main effects)				
Intercept (status at two months)	-0.22(-0.33, -0.12)	-0.23(-0.33, -0.12)	-0.55(-0.63, -0.47)	-0.56(-0.64, -0.48)
Age (linear term)	-0.26(-0.28, -0.24)	-0.26(-0.28, -0.24)	-0.14(-0.15, -0.12)	-0.13(-0.15, -0.12)
Age^2(Quadratic term)	0.008(0.006, 0.009)	0.008(0.006, 0.009)	0.007(0.006, 0.008)	0.007(0.006, 0.008)
Course of CMD				
Both times negative				
Positive only at pregnancy	0.02(-0.22, 0.26)	0.11(-0.20, 0.41)	-0.07(-0.28, 0.14)	-0.07(-0.34, 0.21)
Incident postnatal	-0.31(-0.81, 0.19)	-0.56(-0.19, 0.06)	0.17(-0.27, 0.61)	0.05(-0.52, 0.63)
Persistent CMD	-0.30(-0.74, 0.13)	-0.26(-0.80, 0.29)	0.17(-0.12, 0.56)	--
Fixed effect (Interaction effect)				
Positive only at pregnancy with time		-0.01(-0.04, 0.01)		-0.02(-0.04, 0.003)
Incident postnatal with time		0.04(-0.02, 0.09)		-0.01(-0.05, 0.03)
Persistent CMD with time		-0.006(-0.05, 0.04)		-0.02(-0.05, 0.01)
Variance component				
Level 1 : Within person	1.16(1.13, 1.20)	1.16(1.13, 1.20)	0.90(0.88, 0.93)	0.90(0.88, 0.93)
Level 2 : Status at two months of age	0.97(0.89, 1.07)	0.97(0.89, 1.06)	0.81(0.74, 0.88)	0.81(0.74, 0.88)
Linear term: Standard deviation	0.05(0.03, 0.06)	0.05(0.03, 0.06)	0.04(0.03, 0.05)	0.04(0.03, 0.05)
correlation with at two month status	-0.29(-0.48, -0.09)	-0.29(-0.47, -0.08)	0.04(-0.20, 0.29)	0.05(-0.20, 0.30)
Goodness of fit : Log likelihood	-7644.6	-7643.2	-6724.7	-6722.7
AIC	15309.2	15312.4	13469.5	13471.4
BIC	15373.1	15395.4	13533.5	13554.6

Table 7.25: Effect of course of CMD on the trajectory of logit of underweight and logit of stunting of infant data from the P-MaMiE study

	Course of CMD as predictor of random intercept of logit of stunting *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of logit of stunting* *Estimate (95% CI)*	*Course of CMD as predictor of random intercept o f logit of underweight* *Estimate (95% CI)*	*Course of CMD as predictor of random intercept and random slope of logit of underweight* *Estimate (95% CI)*
Fixed Effects (Main effects)				
Intercept (status at two months)	-2.03(-2.24, -1.83)	-2.02(-2.23, -1.82)	-2.83(-3.28, -2.37)	-2.80(-3.26, -2.35)
Age (linear term)	0.19(0.17, 0.22)	0.19(0.17, 0.22)	0.25(0.16, 0.34)	0.25(0.16, 0.34)
Age^2(Quadratic term)		0.19(0.17, 0.22)	-0.02(-0.02, -0.01)	-0.02(-0.02, -0.01)
Course of CMD				
Both times negative	Ref	Ref	Ref	Ref
Positive only at pregnancy	0.07(-0.34, 0.47)	0.11(-0.14, 0.64)	0.19(-0.37, 0.74)	0.07(-0.61, 0.75)
Incident postnatal	0.28(-0.55, 1.11)	0.27(-0.78, 1.32)	-1.17(-2.51, 0.16)	-1.79(-3.80, 0.23)
Persistent CMD	0.01(-0.72, 0.74)	-0.50(-1.55, 0.55)	-0.30(-1.35, 0.75)	-0.63(-2.00, 0.74)
Fixed effect (Interaction effect)				
Positive only at pregnancy with time		-0.008(-0.07, 0.05)		0.02(-0.05, 0.10)
Incident postnatal with time		0.002(-0.13, 0.13)		0.09(-0.11, 0.29)
persistent CMD with time		0.090(-0.03, 0.21)		0.06(-0.09, 0.21)
Variance component				
Level 1 : Within person				
Level 2 : Status at two months of age	0.77(0.46, 1.30)	0.77(0.46, 1.29)	1.44(1.05, 1.97)	1.44(1.05, 1.97)
Linear term: Standard deviation	0.12(0.09, 0.17)	0.12(0.09, 0.17)	0.17(0.12, 0.24)	0.17(0.12, 0.23)
correlation with at two month status	0.49(-0.55, 0.93)	0.50(-0.56, 0.94)	0.16(-0.33, 0.58)	0.17(-0.33, 0.59)
Goodness of fit : Log likelihood	2473.0	-2471.8	-1905.9	-1905.1
AIC	4962.0	4965.7	3829.8	3834.1
BIC	5013.1	5036.0	3887.4	3910.9

7.5.3 Adjusted effects of perinatal CMD on growth outcomes

7.5.3.1 Adjusted effects of CMD on initial growth of infants – no level-2 predictors of the rate at which growth outcomes change over time

Regression coefficients and their corresponding 95%CI for six growth outcomes and also OR (95%CI) for the two binary growth outcomes are summarized in table 7.26. The figures provided in the body of the table are effect sizes associated with maternal CMD when all variables are included in a fully adjusted model as predictors of initial value of a specific growth outcome. In these fully adjusted MGMs where the growth outcomes are one of length, weight, length-for-age z, weight-for-age z, logit of stunting or logit of underweight, maternal CMD was not a significant predictor of the growth outcome that infants has attained at the age of two months.

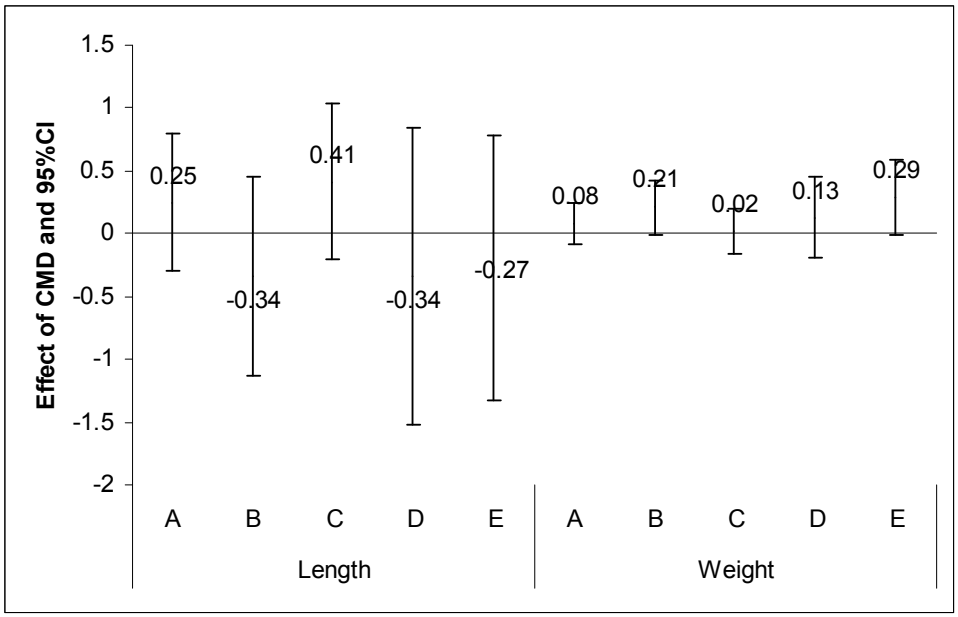

Figure 7.13: Adjusted effect of CMD on length and weight of infants in a model that assumes all covariates as predictors of initial growth but not as predictors of rate of change (A = CMD during pregnancy - prevalence, B = CMD at two month postnatal - prevalence, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

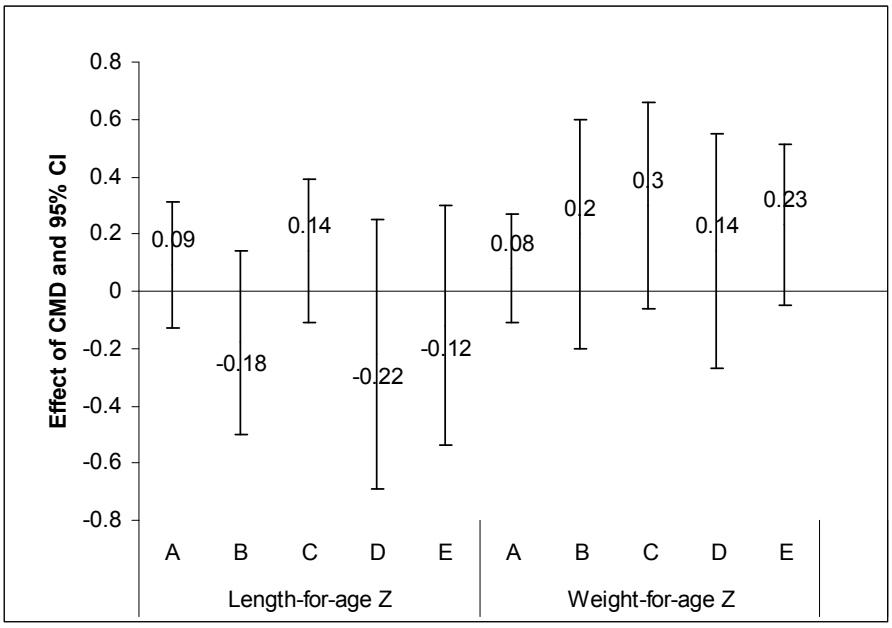

Figure 7.14: adjusted effect of CMD on length-for-age and weight-for-age of infants in a model that assumes all covariates as predictors of initial growth but not as predictors of rate of change (A = CMD during pregnancy, B = CMD at two month postnatal, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

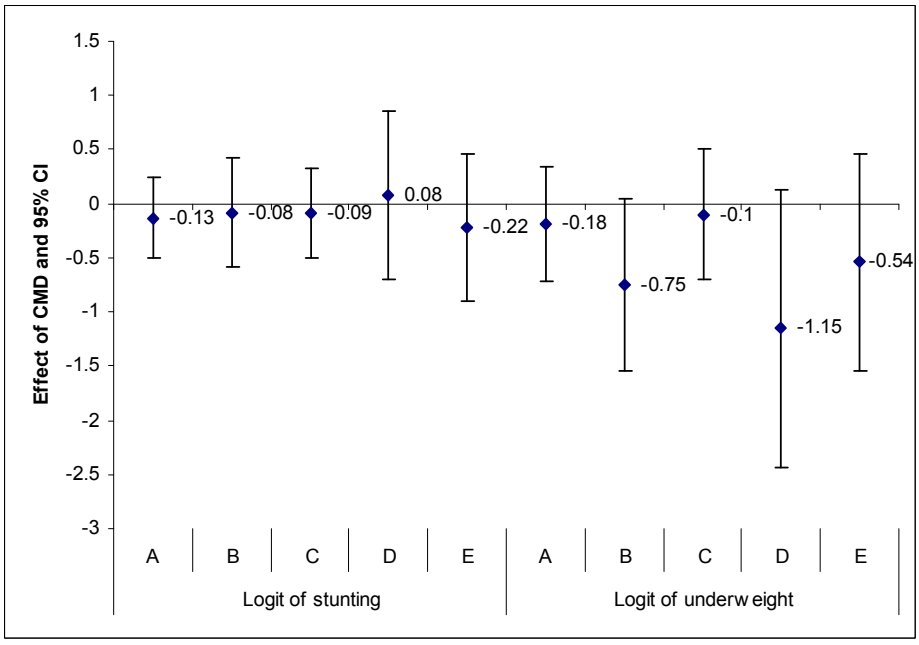

Figure 7.15: adjusted effects of CMD on the logits of stunting and underweight of infants in a model that assumes all covariates as predictors of initial growth but not as predictors of rate of change (A = CMD during pregnancy , B = CMD at two month postnatal, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

Table 7.26: Adjusted effects of CMD on infant growth at initial point (2 months of age) in Butajira Birth cohort, Ethiopia [Interaction terms not included in a model]

Maternal CMD Status at various time points	*Adjusted $\hat{\beta}$ (95% CI)*			
	Length	**Weight**	**Length-for-age**	**Weight-for-age**
Antenatal CMD (prevalent cases)	0.25(-0.30, 0.80)	0.08(-0.08, 0.23)	0.09(-0.13, 0.31)	0.08(-0.11, 0.27)
Postnatal CMD (prevalent cases)	-0.34(-1.14, 0.45)	0.21(-0.01, 0.43)	-0.18(-0.50, 0.14)	0.23(-0.04, 0.50)
Course of CMD				
None at both time	Reference	Reference	Reference	Reference
Antenatal only	0.40(-0.22, 1.02)	0.02(-0.16, 0.19)	0.14(-0.11, 0.39)	0.02(-0.20, 0.23)
Postnatal only	-0.34(-1.51, 0.84)	0.13(-0.20, 0.45)	-0.22(-0.70, 0.25)	0.14(-0.26, 0.55)
Chronic case	-0.27(-1.31, 0.77)	0.28(-0.01, 0.57)	-0.12(-0.54, 0.30)	0.30(-0.05, 0.66)
	Adjusted $\hat{\beta}$ (95% CI) Logit of stunting	**Adjusted $\hat{\beta}$ (95% CI) Logit of underweight**	**OR (95%CI) Stunting**	**OR (95%CI) Underweight**
Antenatal CMD (prevalent cases)	-0.13(-0.49, 0.24)	-0.18(-0.70, 0.35)	0.88(0.61, 1.27)	0.84(0.50, 1.42)
Postnatal CMD (prevalent cases)	-0.08(-0.60, 0.43)	-0.76(-1.56, 0.04))	0.92(0.55, 1.54)	0.47(0.21, 1.04)
Course of CMD				
None at both time	Reference	Reference	Reference	Reference
Antenatal only	-0.09(-0.51, 0.32)	-0.10(-0.69, 0.50)	0.91(0.60, 1.38)	1.11(0.50, 1.65)
Postnatal only	0.08(-0.70, 0.86)	-1.15(-2.44, 0.13)	0.92(0.50, 2.36)	0.32(0.09, 1.14)
Chronic case	-0.22(-0.89, 0.46)	-0.54(-1.53, 0.46)	0.80(0.41, 1.58)	0.58(0.22, 1.58)

7.5.3.2 Adjusted effects of CMD on initial growth of infants – all covariates are included as level-2 predictors of initial value and selected covariates are included as predictors of the rate at which growth outcomes change over time

The adjusted effects of CMD with their corresponding 95%CI on length and weight, on length-for-age z and weight-for-age z, and on logit of stunting and logit of underweight are presented in figures 7.16, 7.17 and 7.18, respectively. Overall there is no significant effect of CMD on growth measures in this group of infants. However, there is a non-significant trend showing that infants whose mothers had experienced perinatal CMD are more likely to be heavier, more likely to attain higher weight-for-age z score and less likely to be underweight at the age of two months. The overall trend with regard to length, length-for-age z and the likelihood of stunting of infants whose mothers are exposed to CMD during pregnancy is also the same but the reverse is true if the exposure is incident postnatal or persistent during perinatal period.

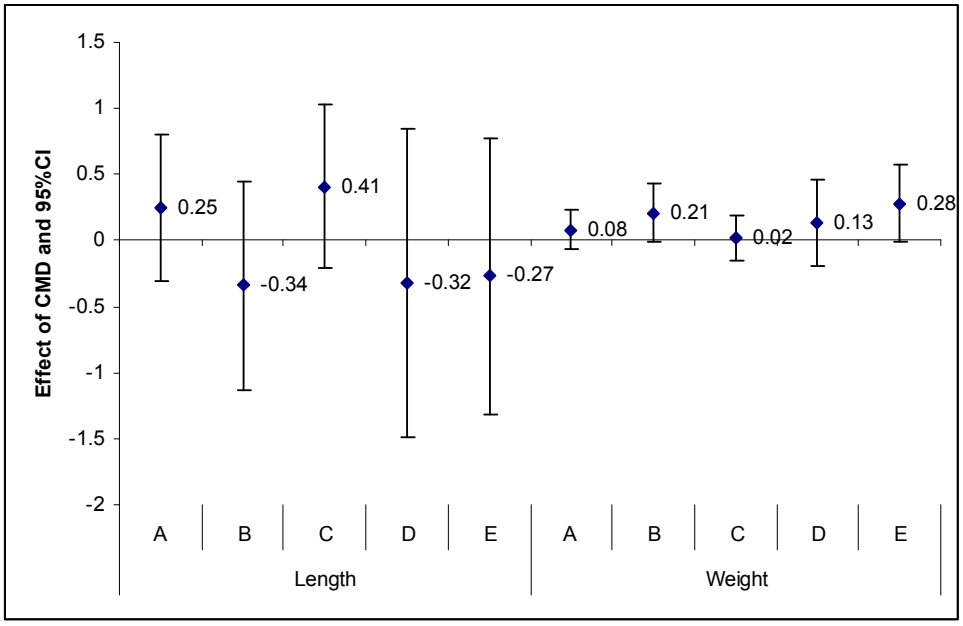

Figure 7.16: Adjusted effects of CMD on length and weight of infants in a model where all covariates are assumed as predictors of initial growth and selected covariates are include as predictors of the rate of growth (A = CMD during pregnancy - prevalence, B = CMD at two month postnatal - prevalence, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

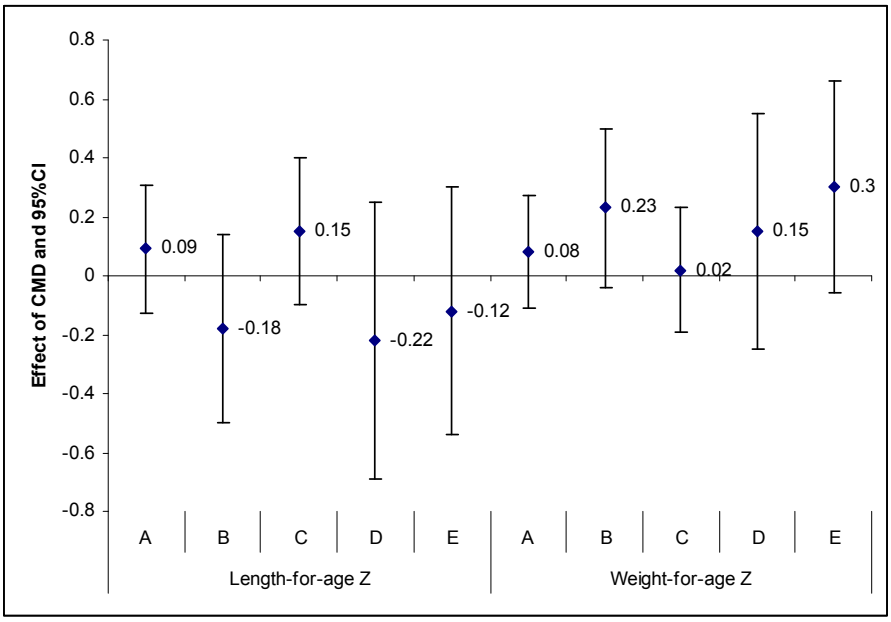

Figure 7.17: Adjusted effects of CMD on length-for-age z and weight-for-age z of infants in a model where all covariates are assumed as predictors of initial growth and selected covariates are include as predictors of the rate of growth (A = CMD during pregnancy, B = CMD at two month postnatal, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

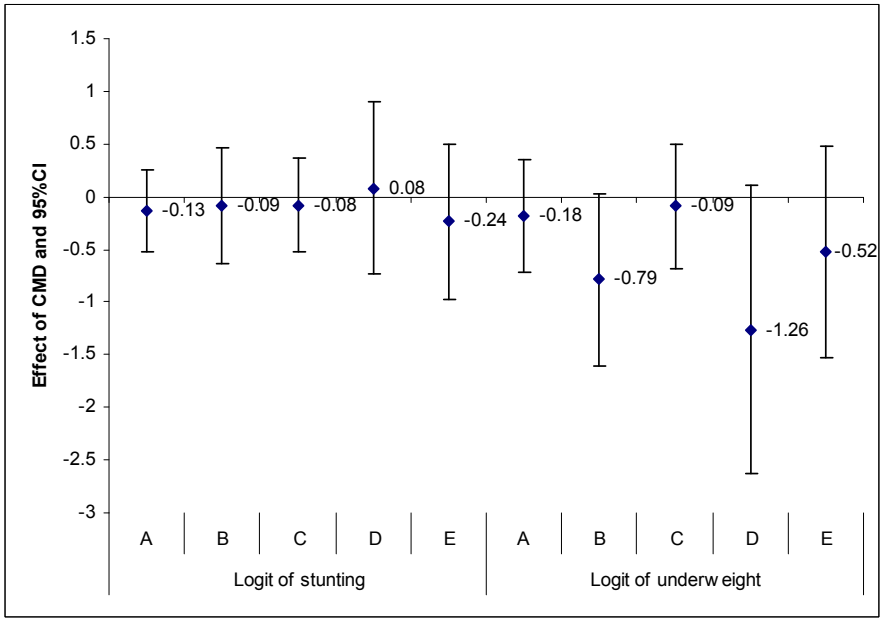

Figure 7.18: Adjusted effects of CMD on logit of stunting and logit of underweight of infants in a model where all covariates are assumed as predictors of initial growth and selected covariates are include as predictors of the rate of growth (A = CMD during pregnancy, B = CMD at two month postnatal, C = CMD at pregnancy which was resolved after birth, D = incident postnatal CMD, E = Chronic CMD that sustained from pregnancy to two month postnatal)

Table 7.27 – Effect of CMD after adjusting for all covariates and interaction terms of selected covariates

Timing for CMD	*Length*	*Weight*	*Length-for-age*	*Weight-for-age*
Third trimester	0.25(-0.30, 0.80)	0.08(-0.08, 0.23)	0.09(-0.13, 0.31)	0.08(-0.11, 0.27)
Two month postnatal	-0.34(-1.13, 0.45)	0.21(-0.01, 0.43)	-0.18(-0.50, 0.14)	0.23(-0.04, 0.50)
Course between the two time points				
None	Ref	Ref	Ref	Ref
Third trimester only	0.41(-0.21, 1.03)	0.02(-0.16, 0.19)	0.15(-0.10, 0.40)	0.02(-0.20, 0.23)
Two month postnatal only	-0.32(-1.49, 0.85)	0.13(-0.20, 0.46)	-0.22(-0.69, 0.25)	0.15(-0.26, 0.55)
At both time points	-0.27(-1.32, 0.77)	0.28(-0.01, 0.57)	-0.12(-0.54, 0.30)	0.30(-0.06, 0.66)
	Logit coefficient **Stunting**	***Logit coefficient*** **Underweight**	OR (95%CI) **Stunting**	OR (95%CI) **underweight**
Third trimester	-0.13(-0.51, 0.26)	-0.18(-0.70, 0.35)	0.88(0.60, 1.30)	0.84(0.49, 1.42)
Two month postnatal	-0.09(-0.65, 0.46)	-0.79(-1.60, 0.02)	0.90(0.52, 1.57)	0.45(0.20, 1.02)
Course between the two time points				
None	Ref	Ref	Ref	Ref
Third trimester only	-0.08(-0.53, 0.36)	-0.09(-0.68, 0.50)	0.92(0.59, 1.44)	0.91(0.51, 1.66)
Two month postnatal only	0.08(-0.74, 0.90)	-1.26(-2.62, 0.10)	1.08(0.48, 2.46)	0.28(0.07, 1.11)
At both time points	-0.24(-0.96, 0.49)	-0.52(-1.51, 0.48)	0.77(0.38, 1.59)	0.60(0.22, 1.61)

CHAPTER 8: DISCUSSION

8.1 Overall summary

8.1.1 Study methodology, strengths and limitations

Study methodology

In this population-based prospective study conducted in a predominantly rural part of Ethiopia we evaluated the effect of maternal CMD in pregnancy and at two months postnatal upon infant growth in the first 18 months of life before and after adjusting for potential risk factors. We also evaluated the effect of pre-specified risk factors on infant growth. Length in centimeters, weight in kilograms, length-for-age z, weight-for-age z, stunting and underweight measured at two, six, nine, twelve and eighteen months of infant age were used as indicators of infant growth. Maternal CMD was measured using a locally validated self reporting questionnaire composed of twenty yes/no items (SRQ-20). We reported three different sets of results generated from (1) traditional modeling techniques (descriptive summary, logistic regression and linear regression at two, at six and at twelve months of infant age), (2) latent growth modeling (LGM) and (3) multilevel growth modeling (MGM). For the LGM a four month time point was created as an additional time point from growth measurements obtained between three and five months of infant age.

Strengths of the study

The strength of this study comes from its longitudinal design, being a large population-based sample from an area with a high prevalence of infant undernutrition and low levels of loss to follow-up over the study period, use of several outcome variables (i.e. binary, standardized continuous and unstandardized continuous variables) to model infant nutritional status and use of different modeling techniques (i.e. cross-sectional and longitudinal modeling techniques) to fully utilize the available information in the data. Unlike most previous studies which were mostly cross-sectional, we were able to measure exposure variables prospectively, free of information bias, and to assess the stability of the association across the neonatal and infant periods of development. This is the first study from sub-Saharan

Africa and the second from a LAMIC setting to ascertain CMD during pregnancy as well as at two months postnatally and to assess their effects on infant growth. We were able to compare the patterns of association with the outcome of interest when analysed as a binary variable (i.e. underweight and stunting) for easy practical interpretation and as a continuous variable (i.e. weight-for-age z score and height-for-age z score) to maximize statistical power. To our knowledge this is the first study to evaluate the effect of CMD on unstandardized weight and height using MGM and LGM. Unlike cross-sectional studies that rely on the recall of parents to determine infant age, which would be difficult in our study setting, we have calculated age based on the contemporaneously recorded date of birth. This information is not routinely available given the low percentage of women in Ethiopia who give birth in clinics or hospitals.

Limitations of the study

A number of limitations can be identified that might have affected our findings. The SRQ-20 is a scale-based measure of maternal CMD symptoms, rather than providing a definitive diagnostic assessment of mental disorder. In three (Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Adewuya, Ola et al. 2008) out of the four (Tomlinson, Cooper et al. 2006) studies that made use of standardised clinical diagnostic measures of maternal depression, a positive association with infant undernutrition was detected. That said, the SRQ-20 has been used extensively in the study area for assessment of CMD in the general population (Alem, Kebede et al. 1999) and was validated before the current study on pregnant and postnatal women from the same geographical area (Hanlon, Medhin et al. 2008). Nevertheless, the assessment of CMD in this setting is by no means straightforward (Hanlon, Medhin et al. 2008) and misclassification of cases is likely to have biased any genuine association towards the null. The low prevalence of maternal CMD that we observed postnatally would also have reduced the study power to detect an effect on infant undernutrition, potentially leading to type II error.

Most of the risk factors were obtained from maternal report as there was no other means of obtaining that information, and as such they could have been subject to recall bias. However,

the outcome was measured prospectively and so maternal recall bias is less likely to have affected the observed associations. Several research workers were involved in measuring height and weight outcomes, which might have introduced measurement errors. Inter-rater reliability was not assessed formally. The training, close follow-up by me and Dr Charlotte Hanlon jointly responsible for field work supervision, periodic quality-control checks and re-measuring of any suspected error cases within one week, should have minimised measurement errors. Any resulting error is likely to be random, and would have had the effect of reducing the size of any genuine effect towards the null. Measuring birth weight only in selected sub-districts might have influenced some of the associations, although there was no obvious reason to assume that birthweight was missing not at random**.** To account for this we included a three category variable (normal, low birth weight, not measured) in our cross-sectional analysis and the effect of low birth weight was in the expected direction, increasing our confidence in our findings. In LGM and MGM missing birthweight was handled using full information maximum likelihood. Although the two month window used to define the outcome could potentially compromise the precision of prevalence estimates of undrnutrition at six and twelve month of age, comparison of the prevalence of underweight and stunting for infants measured within and outside the target one month window did not reveal significant difference. Mortality accounted for the majority of infants lost to follow-up; a higher prevalence of undernutrition within this group might have biased our point estimates. Another limitation of the study is that participating women were not able to correctly report their LMP and so we were unable to ascertain exact gestational age.

8.1.2. Summary of cross-sectional analysis

8.1.2.1 Effect of CMD on infant growth

We investigated the prevalence and trends of infant undernutrition in the first 18 months of infancy and the predictors of anthropometric growth within the first one year of infancy emphasising the role of perinatal CMD. To understand the influence of perinatal CMD and other pre-specified risk factors on growth trajectories we analyzed growth data at three time points and interpreted the consistency of the pattern of association over time. Such studies are rare in resource poor countries given the lack of vital registration. Our finding showed

that the prevalence of infant undernutrition, indicated by stunting (length for age z score less than -2) and being underweight (weight for age z score less than -2) was high starting as early as two months of age and worsened continuously to the age of 18 months. The prevalence of maternal CMD was relatively low, particularly at the two month and at the twelve month postnatal time-points.

CMD during pregnancy which was resolved after birth was significantly associated with better performance in length at two months before and after adjusting for other risk factors. Before adjusting for other risk factors but not in the multivariable analysis prevalent postnatal and incident postnatal CMD were inversely associated with length-for-age z at two month. These associations were also implicated in a non-significant trend of the effect of prevalent and incident postnatal CMD as being risk factors of stunting and for having shorter length at two months of age. At six and twelve months of age there was no significant effect of perinatal CMD on infant length, length-for-age z and stunting.

In bivariate analysis but not after adjusting for other risk factors CMD during pregnancy which was resolved after birth was significantly associated with the risk of underweight at twelve months. Crude and adjusted result also showed non-significant trend in risk of underweight at six months associated with prevalent antenatal CMD and antenatal CMD which was resolved after birth. In bivariate analysis chronic CMD was associated with having increased mean weight at twelve months and in multivariable analysis it was associated with increased weight-for-age z at two and twelve months of age. Postnatal CMD (i.e. prevalent) was significantly associated with an increase in weight-for-age z at two months after adjusting for other risk factors which was not otherwise significant in bivariate analysis. At six months of age perinatal CMD was not significantly associated with underweight and weight-for-age z but there was a positive effect of postnatal CMD (i.e. prevalent) on infant weight which was not significant in multivariable analysis.

8.1.2.2 Risk factors of compromised infant weight other than perinatal CMD

Risk factors of compromised infant weight are dependent on the age of infants and on how weight is entered into the regression model (i.e. weight in kg, weight-for-age z and

underweight). Overall reduced maternal MUAC, male gender, and low birthweight were significant risk factors for compromised infant weight in the first year of infancy but poor sanitary condition of the household and rural residence were significant risk factors of compromised weight after 6 months of age. At two months of age (a) low birthweight, female gender and reduced maternal MUAC were significant predictors of infant weight measured in kg, (b) reduced maternal MUAC, shorter maternal height, higher score on poor sanitary condition scale, male gender and low birthweight were significant predictors of reduced weight-for-age z score, and (c) male gender and low birth weight were significant predictors of underweight. At 6 and 12 months of infant age reduced maternal MUAC, rural residence, higher score on poor sanitary condition scale, male gender, and low birthweight were significant risk factors of compromised weight in its three different forms with the exception that (a) infant gender was not associated with underweight at 12 month, and (b) rural residence was significant risk factor in the adjusted model only for underweight at both time points and for weight-for-age at 6 month.

The negative effect of increased parental age on weight, weight-for-age z and underweight was not observed at two months, it was marginally significant or significant in bivaraite analysis at six months with three weight derived outcomes, and was significant in bivariate or multivariable analysis at twelve months. In bivariate analysis the increased value of poverty index was a significant risk factor of compromised weight, lower weight-for-age and underweight at six and twelve months of age but not at two month. Infants whose mothers had at least one obstetric complication during delivery had a tendency of performing well (a) in their weight, weight-for-age and of not being underweight at two months, (b) on their weight at six months and (c) on their weight and weight-for-age at twelve months.

Early infant feeding practices of the mothers were not significantly associated with infant weight to the extent that they could substantially influence the overall weight gain of infants. In bivariate analysis delayed initiation of breast feeding had a tendency of affecting infant weight at two month of age (weight in kg and weight-for-age z) and at six months of age (weight-for-age z and underweight) and denial of colostrums was negatively associated with weight in kg and underweight at six months and with weight-for-age z at twelve months of

age. In multivariable analysis the only significant result was the negative effect of delayed initiation of breast feeding on weight-for-age z at six months of age.

8.1.2.3 Risk factors of compromised infant length other than perinatal CMD

Similar to that of weight risk factors of compromised infant length are dependent on the age of infants and on how length is used in the model. Overall reduced maternal MUAC, rural residence, poor sanitary condition of the household, male gender and low birth weight were significant risk factors for compromised infant length at 2, 6 and 12 months of age and reduced maternal MUAC was significant predictors of compromised length at 6 and 12 months of age. At two months of infant age (a) maternal MUAC is not significantly associated with infant length, (b) low birthweight is significant predictor of shorter infant length but not associated with length-for-age z and stunting, (c) higher score on poor sanitary condition of the household is significantly associated with increased length, increased length-for-age z and lower risk of stunting, (d) rural residence was significant predictor of shorter length, lower length-for-age z and increased risk of stunting, (e) male gender was significantly associated with increased length, smaller length-for-age z and increased risk of stunting, When infants were six months old (a) only male gender was significant risk factor of stunting and female gender was significant predictor of shorter length and increased length-for-age z, (b) low birthweight was significant predictors of shorter length and smaller length-for-age z, (c) rural residence and reduced maternal MUAC were significantly associated with shorter length, smaller length-for-age z but not with the risk of stunting. At twelve months of age low birthweight, female gender, increased score on poor sanitary scale and rural residence are significantly associated with compromised infant length (short length, lower length-for-age and stunting). The exceptions were that male gender was significantly associated with increased length. An increased MUAC was significantly associated with increased infant length in bivariate analysis but not in multivariable analysis, it is significantly associated with small length-for-age z and not associated with the risk of stunting. In the adjusted model poverty index was not significantly associated with length, length-for-age and stunting at all time points.

Early infant feeding practices of the mothers were not significantly associated with growth outcomes to the extent that they could substantially influence the overall growth of infants. The only significant associations were between (a) denial of colostrums and increased length-for-age at two months in bivaraite and multivariable analysis and (b) denial of pre-lacteal food and smaller length-for-age at two months of age in multivariable analysis but not in bivariate analysis.

8.1.3 Summary of findings from LGM

8.1.3.1 Patterns of infant growth

Change of weight-for-age z score and length-for-age z score within individual were best described by quadratic LGM, and the remaining growth outcomes were best described by non-linear LGM which are not quadratic. In this population a two month old average infant (a) attains a weight of 5.13kg and length of 57.3cm, (b) attains 0.24 and 0.55 standard deviation units below the median score of WHO reference population on its length-for-age z score and weight-for-age z score, respectively, and (c) has a 10.3% probability of being underweight and a 13.5% probability of being stunted. There is significant variability in all initial values of the infants' growth measures except the initial probability of stunting.

Between two and eighteen months of age an average infant gains 4.5kg in weight and 17.5cm in length, and the total gain in both measures varies significantly between infants. Compared to the 2006 WHO child growth standard an average infant in this population attains smaller length-for-age z score and weight-for-age z score during the whole follow-up period. This negative change in length-for-age z score was homogeneous between infants but there was significant variability in case of weight-for-age z score. At the age of eighteen months the probability of stunting of an average infant increased to 62.7%. Probability of underweight reached a maximum level of 21.5% at one year of age and then declined to 18.8% at the age of eighteen months. There was no significant variability between infants in the overall change in the probit of stunting or in the probit of underweight between two and eighteen months of age.

8.1.3.2 The effect of CMD on the patterns of infant growth

In a conditional LGM where antenatal and postnatal CMD were the only predictors of LGM parameters and postnatal CMD was allowed to partially mediate the effect of antenatal CMD on growth parameters there was significant continuity of CMD from antenatal to two month postnatal period. Persistence of CMD has significant negative effect on initial value, positively associated with the linear slope and inversely associated with the quadratic term of length-for-age z. Postnatal CMD was significantly associated with shorter initial value and positively associated with linear slope term of length-for-age z score. Antenatal CMD was significantly associated with increased initial values and inversely associated with the linear slope terms of length-for-age z and weight-for-age z scores. Moreover, antenatal CMD was significantly associated with (a) smaller initial probability of stunting, and (b) smaller overall weight change and higher overall change in the probit of stunting between two and 18 months of age.

In a conditional LGM where birthweight, early infant feeding practices and infant illness are considered as mediators (a) perinatal CMD did not have significant direct effect on length and on the probit of the probability of underweight, (b) there was significant continuity of CMD from antenatal to postnatal period, (c) persistent CMD and prevalent postnatal CMD had significant indirect negative effect on the linear slope factor of length-for-age z through diarrheal episodes, and (d) postnatal CMD was significantly associated with reduced initial value and larger linear slope of length-for-age z score. Antenatal CMD was significantly associated (a) with increased initial values of weight-for-age z and length-for-age z scores, and inversely associated with the linear slopes of weight-for-age z score and length-for-age z score, and (b) with smaller probability of stunting at two months, with higher overall change in the probit of the probability of stunting and with smaller overall change in the weight of infants between two and 18 months. There was no significant indirect effect of antenatal CMD on child growth parameters through birthweight. Although an increase in birthweight was significantly associated with increased initial values of weight, length, length-for-age z, weight-for-age z, lower probability of underweight at two months, and a smaller overall

change in probit of stunting between two and 18 months, there was no significant direct effect of antenatal CMD on birthweight. This implies non-significant indirect effect of antenatal CMD on infant growth through birthweight. Similarly, there were no significant indirect effects of antenatal CMD on child growth parameters through early infant feeding practices.

In a fully adjusted LGM antenatal CMD (a) has significant negative effect on the overall change of weight from two to 18 months, (b) significantly associated with higher initial values of weight-for-age z and length-for-age z, and inversely associated with the linear slope of both growth outcomes, (c) significantly associated with small probability of stunting at initial time point and (d) it is not significantly associated with birthweight implying absence of its indirect effect on growth through birthweight. Postnatal CMD is significantly associated with smaller probability of stunting at initial time and positively associated with the overall change in the probit of the probability of stunting between two and 18 months of age. There is significant continuity of CMD from antenatal to postnatal period and this persistent CMD is significantly associated with increased probability of stunting at initial time point and significantly associated with smaller overall change in the probit of the probability of stunting over the study period. Indirect effect of persistent CMD on the overall change of the probit of stunting is marginally non-significant. However, there is significant indirect negative effect of persistent CMD on the linear slope of length-for-age through diarrheal episodes during early age of infancy.

8.1.3.3 Factors other than CMD associated with infant length

Higher score on poor sanitary condition scale and larger birthweight are significantly associated with increased initial length and increased initial values of length-for-age z. Similarly, being a female is significantly associated with increased initial value of length-for-age z. increased maternal age and not having any obstetric complication were significantly associated with small overall change of length during the study period and smaller birthweight is significantly associated with increased overall change in length-for-age z during the study period. Smaller score on poor sanitary condition scale is associated

with larger overall increases in length and length-for-age z score. Being female is significantly associated with increased overall change in length-for-age and small overall change in length over the study period. Urban residence, female gender and increased birthweight are significantly associated with reduced initial probability of stunting and lower score in poor sanitary condition scale and delayed initiation of breast feeding are significantly associated with small overall change in the probit of the probability of stunting during the study period

8.1.3.4 Factors other than CMD associated with infant weight

Increased maternal height, increased maternal MUAC, having at least one obstetric complication, increased birthweight, and being male infant are significant predictors of increased initial weight. Similarly, increased maternal height, being a female infant and increased birthweight are significant predictors of increased initial value of weight-for-age z score. Increased maternal age and higher score on poor sanitary condition scale are significant predictors of smaller overall change in weight and smaller linear change in weight-for-age z during the study period. Moreover, smaller birthweight is significantly associated with larger linear slope of weight-for-age z. None of the risk factors considered in this analysis was significantly associated with the probit of the overall change of the probability of stunting. However, being a female infant, increased birthweight, lower score on poor sanitary condition scale, increased maternal height, having at least one obstetric complication and increased material age are significant predictors of smaller probability of underweight at the age of two months.

8.1.4 Summary of findings from MGM

8.1.4.1 Patterns of infant growth

Four candidate polynomial functions (i.e. random intercept but no growth over time, random intercept and non-random linear growth, random intercept and linear growth with random slopes and quadratic growth over time with random intercepts and random slopes) were evaluated to establish the best fitting function which describes infant growth (i.e. length, weight, length-for-age z, weight-for-age z, logit of stunting and logit of underweight) between two and eighteen months of age. The time of growth measurements were centered at two months to give the interpretation of initial value for the intercept term.

Except logit of stunting which was best described by a linear function with random intercepts and linear slope all the growth outcomes were best described by a quadratic function with random intercepts and random linear terms. At the age of two months there was significant variability in growth levels of individual infants. Initial length-for-age z and weight-for-age z are significantly and inversely correlated with linear slope of an average infant. During the follow-up period there was significant change in all growth outcomes of an average infant (i.e. an increase in length, weight, logit of stunting and logit of underweight, and a decrease in length-for-age z and weight-for-age z) with significant variability between individual infants. The average infant's initial weight, length, logit of stunting and logit of underweight were not significantly correlated with the linear slope. Between two and eighteen months of age there was a monotonic increase in weight and length of an average infant and a monotonic decline in length-for-age z and weight-for-age z. During the follow-up period (a) there was no significant risk of underweight of an average infant and (b) the risk of stunting was statistically significant at the age of 18 months but not during the first year of infancy.

8.1.4.2 The effect of CMD on the patterns of infant growth

The overall fit of the best fitting unconditional MGM was significantly improved when (a) antenatal CMD was considered as the only predictor of initial logit of stunting and initial logit of underweight, (b) postnatal CMD was considered as the only predictor of initial length, weight, length-for-age z and weight-for-age z , and (c) the course of perinatal CMD was considered as the only predictor of initial length, weight, length-for-age z, weight-for-

age z and logit of stunting. Relaxing model assumption and allowing maternal CMD to be the only predictor of both initial value and the rate of change of each growth outcome did not result in a significant improvement of the best fitting baseline unconditional MGM.

Before adjusting for other risk factors there was no significant effect of perinatal CMD (i.e. prevalent antenatal, prevalent postnatal or the course of perinatal CMD from pregnancy to postnatal period) on initial values of the growth measures and on the rate at which infants grow over time. The only exception was that infants whose mothers had experienced chronic CMD were significantly heavier at the age of two months compared to infants whose mothers had never experienced perinatal CMD. There was significant within individual growth variability in all the models. Similarly, there was significant variability between infants' initial growth measures as well as in the rate at which infants grow over time. The amount of variability in initial value and in the rate of growth explained by antenatal CMD, postnatal CMD or the course of perinatal CMD was extremely low.

In a fully adjusted MGM where infant growth outcome is one of length, weight, length-for-age z, weight-for-age z, logit of stunting or logit of underweight, maternal CMD was not a significant predictor of the growth level attained at the age of two months. This finding was consistent whether predictors of the rate of change were include in or excluded from the MGM. However, there was a non-significant trend showing that two month old infants whose mothers had experienced perinatal CMD are more likely to be heavier, more likely to attain higher weight-for-age z score and less likely to be underweight. Infants whose mothers were exposed to CMD during pregnancy have also similar overall trend with regard to length, length-for-age z and the likelihood of stunting. However, the reverse was true if the exposure was incident postnatal or persistent from pregnancy to two month postnatal.

8.1.4.3 Factors other than CMD associated with infant length

Before adjusting for other risk factors (a) rural residence, female gender, low birth weight, and receiving colostrum were significantly associated with shorter initial length of an average infant, (b) rural residence, male gender and receiving colostrum at birth were

significantly associated with shorter initial length-for-age z of average infant, and (c) being a female was significantly associated with reduced risk of stunting of a two month old average infant. Similarly, (a) increased parental age, scoring higher on poor sanitary condition scale, scoring higher on poverty scale, rural residence and compromised maternal autonomy were inversely and significantly associated with the rate that infants gain their length or length-for-age z, and (b) increased maternal age, scoring higher on poor sanitary condition scale, scoring higher on poverty scale, urban residence and low birth weight were significantly associated with larger rate with which logit of stunting increases.

In a fully adjusted conditional MGM being a male infant, urban residence, increased maternal MUAC and scoring higher on poor sanitary condition scale were significant predictors of increased length and length-for-age z, and reduced logit of stunting of a two month old average infant. However, the statistical significance of the effect of poor sanitary condition scale was marginal. Low birth weight was significantly associated with shorter length but not significantly associated with length-for-age z and logit of stunting of a two month old average infant. Compromised maternal autonomy and denial of colostrum were significantly associated with increased length and length-for-age z attained by a two month old average infant but not with logit of stunting. Similarly, not getting pre-lacteal feeding and being immunized before two months of age were significantly associated with shorter length-for-age z score and increased risk of stunting of a two month old average infant, respectively. Maternal age marginally affected the rate of change of length and logit of stunting. Poor sanitary condition scale was significant predictor of the rate of change of length, length-for-age z and logit of stunting. Maternal autonomy and birth weight were significant predictors of the rate of change of length and logit of stunting, respectively.

8.1.4.4 Factors other than CMD associated with infant weight

Before adjusting for other risk factors (a) increased maternal height, increased maternal MUAC, having at least one obstetric complication , urban residence, having younger father, scoring lower on poor sanitary condition scale, scoring lower on poverty scale, male gender, normal birth weight, and immediate initiation of breastfeeding were significantly associated

with increased average infant's initial weight, (b) increased maternal MUAC, having at least one obstetric complication, urban residence, scoring lower on poor sanitary condition scale, scoring lower on poverty scale, female gender, normal birth weight, and immediate initiation of breastfeeding were significantly associated with increased average infant's initial weight-for-age z, and (c) having at least one obstetric complication , scoring lower on poor sanitary condition scale, being a female infant, and normal birth weight were significantly associated with reduced risk of average infant's initial underweight. Similarly, (a) increased parental age, having at least one previous under five child, scoring higher on poor sanitary condition scale, scoring higher on poverty scale and rural residence were inversely and significantly associated with the rate that infants gain their length or length-for-age z, (b) increased parental age, reduced maternal MUAC, having at least one previous under five child , scoring higher on poor sanitary condition scale and scoring higher on poverty scale were significantly associated with larger rate with which logit of underweight changes, and (c) urban residence was significantly associated with larger rate with which length of an average infant increase and smaller rate with which logit of underweight changes.

In a fully adjusted model (a) not having obstetric complication during delivery, being female infant, low birth weight, scoring higher on a poor sanitary condition scale and delayed initiation of breastfeeding were significant predictors of lighter initial weight, (b) increased maternal MUAC, having obstetric complication during delivery, being a female infant, being born with normal birth weight and scoring lower on a poor sanitary condition scale and immediate initiation of breastfeeding were significant predictors of better initial weight-for-age z, (c) being male infant and being born with low birth weight were significantly associated with increased risk of underweight of a two month old average infant. Scoring higher on poor sanitary condition scale and smaller maternal MUAC were marginally associated with increased initial risk of underweight. Scoring higher on poor sanitary condition scale was significantly associated with smaller rate of change of weight and increased maternal age was significantly associated with smaller rate of change weight and weight-for-age z. However, none of the risk factors considered in the analysis were significantly associated with the logit of underweight.

8.2 Growth levels and prevalence of infant undernutrition in Butajira

In our study area a two month old average infant attains a weight of 5.13kg, a length of 57.3cm, and 0.55 standard deviation units and 0.24 standard deviation units below the median weight and length, respectively, of the 2006 WHO reference standard with significant variability between infants' initial growth measures. The later two figures are larger than the average deviations that would be expected at a global and regional levels (Shrimpton, Victora et al. 2001) implying unfavorable circumstances of our study participants to attain their growth potential at the age of two months. Between two and 18 months of age an average infant in the current study gains weight and length progressively and attains a weight of 9.6 kg and a length of 74.7cm at the age of 18 month. At different ages these infants attained significantly smaller weight compared to infants in Peru (Wachs, Creed-Kanashiro et al. 2005), larger weight compared to infants in India (Kapur, Sharma et al. 2005), shorter length compared to infants in Peru (Wachs, Creed-Kanashiro et al. 2005) and comparable growth levels with infants in Nigeria (Adewuya, Ola et al. 2008). Although we do not have sufficient data to substantiate our arguments differences in a range of factors including genetic, environmental, economic, maternal health and child health might explain the differential growth attainments in the three counties. Our argument is in line with the fact that the counties are from regions with different levels of child growth faltering (Shrimpton, Victora et al. 2001). We are unaware of any published data on the expected length or weight of Ethiopian infants within our target age range limiting our ability to compare growth performance of these infants with the national or local figures. The rates of increase in infants' weight and length in the current study are smaller than would be predicted from the 2006 WHO reference standards. This is evident from a monotonically declining trend observed in our study infants' length-for-age z between two and 18 months of age and weight-for-age z between two and 12 months of age. The monotonic declining trend is consistent with the trend of the nation as a whole (Central Statistical Agency and ORC Macro 2006), other LAMIC (Kikafunda, Walker et al. 1998; Ãsling-Monemi, Naved et al.

2009) and the whole regions of Africa, Asia and Latin America (Shrimpton, Victora et al. 2001).

The prevalence of undernutrition progressively increases in Ethiopia from early infancy, peaking in the late months of the second year of age and thereafter starting to stabilise. The prevalence of stunting starts to exceed that of underweight when undernutrition starts to stabilize (Central Statistical Agency and ORC Macro 2006). The current finding follows the same increasing trend but the prevalence of stunting overtook that of underweight as early as two months of age. This difference is not unexpected in light of the varying degree of undernutrition across different geographical regions of the country (Getahun, Urga et al. 2001; Central Statistical Agency and ORC Macro 2006). However, for 48% and 62.8% of infants to be stunted at one year and at 18 months of age, respectively, while 99.6% and 96.1% of them are still breastfeeding at 12 months and at 18 months, respectively, is of concern. The national prevalence of stunting is 32.7% and 46.3% among 9-11 months old and 12-17 months old, respectively (Central Statistical Agency and ORC Macro 2006). In Burkina Faso(Thiombiano-Coulibaly, Rocquelin et al. 2004) the prevalence of stunting at the mean age of 5 months is less than 2.0% and in a semi-urban population in Uganda the prevalence of stunting among 0-11 month old infants is 16.7% (Engebretsen, Tylleskar et al. 2008). A prevalence of stunting comparable to the current study was reported among 0-12 months old in India (Kumar, Goel et al. 2006). One possible explanation for the high prevalence of stunting at one year and at 18 months could be due to the late introduction of supplementary food with low nutritional quality (Getahun, Urga et al. 2001). Similarities and differences across different studies could be attributed to the underlying socioeconomic and socio-cultural conditions. The Meskan, Mareko and Silti districts to which our Butajira study site belongs have been subject to drought in recent years. During the study period, bi-annual screening for severely undernourished infants and pregnant/lactating women was being conducted in order to provide food supplements. The same screening practice also existed in the study area before the launch of the current study (Teferi, Lera et al. 2010). The high prevalence of stunting is therefore not unexpected.

8.3 The effect of perinatal CMD on infant growth in Butajira

The possible association between maternal CMD and child undernutrition in LAMIC has captured the attention of researchers in recent years, and has been tested using epidemiological studies of varying methodological quality, different measures of child undernutrition, and various measures of CMD that may have contributed to the different findings across settings. However, consistent and significant associations have been observed in south Asia independent of these and other heterogeneities. Recent meta-analysis has reported significant negative effect maternal depression on child underweight and stunting (Surkan, Kennedy et al. 2011)

To our knowledge only one of the previous studies has investigated the effect of CMD on child growth using unstandardized measures of length and weight, and in the adjusted models the negative effect of CMD was not statistically significant (Hazarika 2010). While investigating the effect of CMD on child nutritional status in LAMIC using standardized growth outcomes previous researchers have not used LGM and MGM as their modeling techniques. In our study the effect of CMD on infant length and weight was dependent on the modeling technique used and on how SRQ-20 score was treated in the model. In a cross-sectional analysis exposure to CMD during pregnancy which was resolved after birth compared with no exposure to perinatal CMD was significantly associated with increased initial infant length. This association could possibly be explained by reverse causality. In a bivariate analysis but not in multivariable analysis postnatal CMD or persistent CMD compared to no perinatal CMD were significantly associated with increased weight at the age of six and 12 months, respectively. In LGM higher antenatal SRQ-20 score was significantly associated with smaller weight gain between two and 18 months. This association was not mediated through birth weight, early infant feeding practices, infant illnesses or postnatal CMD. In all other scenario there was no significant direct or indirect effect of CMD on initial weight or length and on the overall change in weight or length during the follow-up period. Our significant finding contrasts with the previous studies which have demonstrated significant negative effect of CMD on length but not on weight (Harpham, Huttly et al. 2005; Stewart, Umar et al. 2008; Black, Baqui et al. 2009) although direct comparison is limited by the difference in the study design, modeling techniques used

and the scale of growth outcome measures (i.e. unstandardized versus standardized growth measures). In MGM there was non-significant trend indicating that perinatal CMD might be associated with better initial weight, antenatal CMD might be associated with better initial length and postnatal or persistent CMD might be associated with compromised initial length. In this socio-economically deprived and rurally dominated community the negative effects of less frequent disorders including CMD on weight and length might have been diluted by the negative effects of more frequent competing risk factors that affect infant growth.

The two previously published population-based cohort studies (Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006), both using diagnostic measures of maternal depression and employing traditional modeling techniques, present conflicting results: in periurban South Africa no association was found with weight-for-age z, length-for-age z, stunting and underweight of index child at both two and 18 months (Tomlinson, Cooper et al. 2006), whereas in rural Pakistan (Rahman, Iqbal et al. 2004) the association was seen with categorical indicators of under-nutrition at both six and 12 months (underweight: OR = 3.5; 95% CI: 1.5 - 8.6 at six months and OR = 3.0; 95% CI: 1.5 - 6.0 at 12 months, and stunted: OR = 3.2; 95% CI: 1.1 - 9.9 at six months; OR =2.8; 95% CI: 1.3 - 6.1 at 12 months). Our study sample is most comparable to the Pakistan study, although socioeconomic measures indicate greater poverty in the Ethiopia sample, for example, substantially lower levels of household electricity and flush toilets compared to Pakistan (Rahman, Iqbal et al. 2004). Contrasting the South African study (Tomlinson, Cooper et al. 2006) and in line with the cross-sectional study in Brazil (Surkan, Kawachi et al. 2008) postnatal CMD in the cross-sectional analysis was significantly associated with larger weight-for-age z score at the age of two months. In LGM postnatal CMD did not have significant direct effect on weight-for-age z or on the probit of underweight. However, higher postnatal SRQ-20 score was significantly associated with (a) a smaller initial length-for-age z score contrasting to the USA study in which postnatal depression was association with increased length (Ertel, Koenen et al. 2010) and the non-significant effect in the South African study (Tomlinson, Cooper et al. 2006), and larger linear change in length-for-age z, and (b) higher risk of stunting at the age of two months by contrast with the South African study (Tomlinson, Cooper et al. 2006) and smaller change in the probit of stunting between two and 18 months

of age. Higher antenatal SRQ-20 score was significantly associated with (a) increased initial length-for-age z contrary to the non-significant effect in the USA study (Ertel, Koenen et al. 2010), increased initial weight-for-age z in contrast to the children in the USA who have started being small and remained smaller at the age of three years (Ertel, Koenen et al. 2010), and reduced initial risk of stunting, and (b) smaller rates of linear changes of length-for-age z and weight-for-age z. Most of our significant findings are counter intuitive and in contrast to the findings in the USA (Ertel, Koenen et al. 2010; Ertel, Koenen et al. 2010). We are unaware of previous studies in LAMIC that have evaluated the effect of CMD on child growth using LGM limiting our ability to compare the findings. Outside of South Asia, most of the negative findings from South America (Harpham, Huttly et al. 2005; Surkan, Ryan et al. 2007) and sub-Saharan Africa (Harpham, Huttly et al. 2005; Tomlinson, Cooper et al. 2006) originated from population-based studies, while most of the positive findings (de Miranda, Turecki et al. 1996; Adewuya, Ola et al. 2008; Stewart, Umar et al. 2008) are from clinic-based studies. The nature of the selection bias is not immediately evident, but the potential is clearly present given the limited access to and use of routine antenatal and obstetric care, particularly in sub-Saharan Africa. The type of study design used to collect the data and the modeling technique used to evaluate the association could be among the factors that might explain the negative findings given the discrepancies observed in our study due to varying outcome measures and use of different modeling techniques.

The timing of measurement of infant undernutrition could have relevance, with the two previously negative studies from sub-Saharan Africa evaluating children at an older age: 18 months (Tomlinson, Cooper et al. 2006) and 6 to 18 months (>50% over 12 months of age) (Harpham, Huttly et al. 2005). Similarly, for the negative study from Jamaica (9 to 30 months) (Baker-Henningham, Powell et al. 2003). In the Nigeria study, a significant association between postnatal CMD and infant undernutrition was only found at three and six, but not at nine months of age (Adewuya, Ola et al. 2008). Although the Bangladesh study found the reverse, that maternal CMD was only associated with infant undernutrition at 12 months and not at six months, this is likely to have occurred because maternal CMD was measured at 12 months and thus showed a stronger association concurrently (Black, Baqui et al. 2009). In our cross-sectional analysis no association with infant underweight and

stunting was apparent at two, six or 12 months of age. Although similar association was reported in Brazil among a cross-sectional sample of children aged 6-24 months (Surkan, Kawachi et al. 2008) reverse causality might not be ruled out from being a plausible explanation for the observed positive effect of postnatal CMD on weight-for-age z at two months and that of persistent CMD on weight-for-age z at two and 12 month. Pregnancy weight gain might be considered as the sign of healthy growing fetus because it is positively correlated with birthweight (Butte, Ellis et al. 2003; Anderson, Bignell et al. 2009). However, it could also predispose pregnant women living in a socio-economically disadvantaged environment like ours to excessive worry about their home delivery. Moreover, delivery of heavier babies could potentially increase the risk of obstetric complication during home delivery followed by an increased demand for breast feeding and infant care. All these facts might expose mothers to incident or persistent CMD explaining the positive association observed in the current study. When repeated growth measurements were analyzed using LGM antenatal CMD was significantly associated with better initial weight-for-age z and length-for-age z, smaller rate of linear growth of weight-for-age z and length-for-age z, and reduced risk of initial stunting. These associations were independent of the effect of persistent CMD on child growth and they were not meditated through birthweight or early feeding practices. Since there was no significant correlation between initial growth and rate of linear growth it is not obvious how antenatal CMD could have a direct positive effect on initial growth and negative effect on the rates of linear growth afterwards. Postnatal CMD was significantly associated with compromised initial length-for-age z, increased risk of stunting at the age of two month, larger rate of linear change in length-for-age z, and smaller change in the level of stunting between two and 18 months of age. Moreover, it was significantly associated with smaller rate of linear change of length-for-age z through its positive effect on the occurrence of diarrheal episodes. Our finding strengths previously observed negative effect of postnatal CMD on infant length (Patel, DeSouza et al. 2003; Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Black, Baqui et al. 2009) and extends the existing knowledge about the association of CMD, diarrhoea and infant growth (Rahman, Iqbal et al. 2004; Rahman, Bunn et al. 2007) by shedding light on one of the possible mechanism through which persistent and postnatal CMD might affect infant length.

Most previous studies reported categorical indicators of infant nutritional status. Where the analyses were repeated for both categorical and continuous nutritional indices, only the categorical measure was associated with maternal CMD in Brazil (Surkan, Ryan et al. 2007; Surkan, Kawachi et al. 2008), and neither were associated in South Africa (Tomlinson, Cooper et al. 2006). When the two have been compared in the same study, impaired linear growth (length-for-age; stunting) has more often been associated with maternal CMD than the composite nutritional indicator of weight-for-age(Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Black, Baqui et al. 2009). In our study we used both continuous and categorical indicators of nutritional status. Contrary to our expectation positive association was observed in a cross-sectional analysis between (a) CMD during pregnancy which was resolved after birth and infant length at two months, (b) prevalent postnatal CMD and weight-for-age z at two months, and (c) persistent CMD and weight-for-age z at two and 12 months. In LGM larger antenatal SRQ-20 score was significantly associated with increased initial values and smaller linear changes of weight-for-age z and length-for-age z, and reduced initial risk of stunting. However, larger postnatal SRQ-20 score was significantly associated with smaller initial value and larger linear change of length-for-age z, increased initial risk of stunting and smaller change in the logit of stunting between two and 18 month of age. The later finding is partly consistent with previous findings (Surkan, Ryan et al. 2007; Stewart, Umar et al. 2008; Black, Baqui et al. 2009) although complete comparison of the result is limited by the absence of previous studies which have used similar study design and modeling techniques.

Persistent perinatal CMD could impact on nutrition during pregnancy and after birth. The nature of any interaction between CMD in pregnancy and the postnatal period to cause under-nutrition is unclear. Contrary to the current findings of cross-sectional analysis there is strong evidence in Pakistan (Rahman, Iqbal et al. 2004) showing that chronic perinatal CMD significantly increases the risk of infant underweight (relative risk (RR) = 5.9; 95% CI: 2.7 to 12.8 at six months, RR = 3.5; 95% CI: 2.2 to 5.6 at 12 months) and stunting (RR = 5.5; 95% CI: 1.9 to 16.0 at 6 months and RR = 3.2; 95% CI: 1.9 to 5.4 at 12 months). However, in the Pakistan study there was little remission of depression in pregnancy, or incidence of postnatal depression. In our cross-sectional analysis a very low prevalence of persistent

CMD compromised the power to detect any meaningful effect on binary growth outcomes. Reverse causality might be considered as one possible explanation for significant positive association between persistent CMD and weight-for-age z at two and 12 months of age. In LGM persistent CMD had significant (a) indirect negative effect on the linear slope of length-for-age z through diarrheal episodes, (b) positive effect on initial probit of stunting, and (c) negative effect on the total change in the probit of stunting between two and 18 months of age. This finding is partly in agreement with Pakistan study (Rahman, Iqbal et al. 2004; Rahman, Bunn et al. 2007) in that CMD affects infant length directly and through diarrheal episodes although direct comparison of the current study and previous studies are limited due to difference in their data analysis methods.

Five previous studies, from Ethiopia (Harpham, Huttly et al. 2005), South Africa (Tomlinson, Cooper et al. 2006), Brazil (Surkan, Ryan et al. 2007), Peru (Harpham, Huttly et al. 2005) and Jamaica (Baker-Henningham, Powell et al. 2003) have failed to replicate the association between maternal CMD and infant undernutrition seen in South Asia (Patel, DeSouza et al. 2003; Anoop, Saravanan et al. 2004; Rahman, Iqbal et al. 2004; Rahman, Lovel et al. 2004; Harpham, Huttly et al. 2005; Black, Baqui et al. 2009). We have previously found that maternal CMD in pregnancy in this Ethiopian cohort (i.e. the same cohort on which this thesis is based on) was not associated with low birth weight (Hanlon, Medhin et al. 2009) which is also true in the current study, again in contrast to the findings from South Asia. Although some positive findings observed in our LGM are indicative for the importance of conducting well designed and powered studies the true absence of an adverse effect of maternal CMD in pregnancy or the postnatal period on child nutritional status in Ethiopia is a possibility. When interpreting their negative findings from Ethiopia and Peru compared to India and Vietnam, Harpham et al. call for qualitative exploration for the reasons for such differences and speculate that the 'pressurised cultural role of women in relation to childcare' in South Asia might be exacerbated by a child who is failing to thrive, leading to worsening maternal mental health (Harpham, Huttly et al. 2005). We might also now add a recommendation for the need of a well powered prospective study which employs a modeling technique like ours that uses the data more efficiently. In Ethiopia, shared parenting practices within families and neighborhoods may have diluted any negative effect of postnatal and persistent CMD that would have been apparent with the binary growth

outcomes. Informal feedback from our project data collectors suggests that children in this community are considered as potential future capital, giving higher parity mothers an elevated social rank compared to mothers of the same age with a smaller number of children. One of the common reasons to justify polygamous marriage in the community is the demand for more children by the husband. This could mean that maternal CMD becomes less prevalent and/or severe as the family expands, and the presence of more siblings for the child also facilitates shared parenting. However, at this stage there are no clear answers for why maternal mental disturbance appears to have such a significant effect on child growth in some countries and not in other countries.

8.4 Risk factors of infant growth other than CMD

Like other African countries (WHO Collaborative Study Team on the Role of Breastfeeding on the Prevention of Infant Mortality 2000) breastfeeding at one year of age is a norm in Ethiopia (Central Statistical Agency and ORC Macro 2006) and in the current study 100%, 99.6% and 96.1% of the infants were breastfeeding at two, 12 and 18 months follow-up, respectively. However, adherence to the optimal breastfeeding practices (WHO 1989) that could reduce infant morbidity and enhance growth (Islam, Ahmed et al. 2006; Diallo, Bell et al. 2009) is still low in Ethiopia (Alemayehu, Haidar et al. 2009). In the current study delayed initiation of breast feeding was significantly associated with compromised initial weight and weight-for-age, larger rate of linear change in length-for-age and smaller change in the probit of stunting between two and 18 months. Receiving prelacteal feed was significantly associated with shorter initial length-for-age but contrary to our expectation receiving colostrum was significantly associated with shorter initial length and smaller initial length-for-age z. Several investigators (Brennan, McDonald et al. 2004; Kumar, Goel et al. 2006; Engebretsen, Tylleskar et al. 2008) have reported an increased risk of infant undernutrition resulting from sub-optimal feeding during early infancy (discarding of colostrum, delayed initiation of breast feeding, pre-lacteal feeding and non-exclusive breastfeeding) and less food intake in early childhood (Kapur, Sharma et al. 2005). However, other investigators did not replicate these findings (Sanghvi, Thankappan et al. 2001; Thiombiano-Coulibaly, Rocquelin et al. 2004). In Bangladesh, food secure households were

more likely to practice sub-optimal infant feeding at the age of 3-6 months (Saha, Frongillo et al. 2008) but the prevalence of undernutrition was significantly lower in this group compared to food insecure households (Saha, Frongillo et al. 2009). In Egypt pre-lacteal feeding was associated with an increased risk of diarrhea and early introduction of supplementary food (Hossain, Radwan et al. 1992). In the current study, the proportion of women delaying initiation of breastfeeding for more than one hour is comparable to the national figure but the proportion of non-excusive breastfeeding at the age of two months and pre-lacteal feeding practices are relatively low (Central Statistical Agency and ORC Macro 2006; Alemayehu, Haidar et al. 2009). The proportion of infants who were given colostrum in the current study is higher than studies from Burkina Faso (Thiombiano-Coulibaly, Rocquelin et al. 2004) and India (Kumar, Goel et al. 2006). One possible explanation for the absence of significant effects of sub-optimum infant feeding practices on stunting and underweight might be a lack of power to detect an effect due to the low prevalence of sub-optimal feeding practices. Compromised initial length is the result of cumulative effects of prenatal and early postnatal factors and receiving colostrum is more likely to contribute positively to the immune system of the infant. Hence, reverse causality might be more plausible explanation for the counter intuitive negative effect of receiving colostrum on initial length-for-age. Since 89% of the study participants were delivered at home and attended by their mothers/ mother-in-low or elderly women living in the neighborhoods, there is a greater likelihood of a disadvantaged infant to get more focus which includes receiving colostrum.

In the current study larger maternal height and larger MUAC in pregnancy were significantly associated with heavier initial weight, larger initial weight-for-age and reduced risk of initial underweight. Similarly, larger maternal MUAC was significantly associated with larger initial length, larger initial length-for-age, reduced initial risk of stunning and with larger linear change of length-for-age. An increase in maternal age was significantly associated with smaller total change in weight and length between two and 18 months, with smaller linear change of weight-for-age, and with larger linear change in the logit of stunting. In a cross-sectional sample of infants aged 5-11 months in two rural villages of Ethiopia, maternal height, triceps skin fold thickness and zinc concentration in breast milk were

associated with stunting but current weight and MUAC were not associated with stunting (Umeta, West et al. 2003). Shorter maternal height and smaller weight gain during pregnancy have previously been found to be significant predictors of severe stunting of infants at one year of age (Espo, Kulmala et al. 2002). In the current study we do not have data to investigate either the effect of weight gain during pregnancy or the effect of the content of breast milk. Stunting is thought to result from both the nutritional experience of the individual over a period of time, as well as the nutritional status of the parents, particularly the mother, with the maternal effect mainly acting through birth weight (Ramakrishnan, Martorell et al. 1999) and possibly birth length (Schmidt, Muslimatun et al. 2002). The current finding is independent of the confounding and mediating effects of birthweight. However, we were not able to adjust for the possible mediating effect of birth length. The positive effects of maternal height and MUAC on infant's initial growth levels and the positive effect of MUAC on linear change of length-for-age suggests that in this setting interventions that improve the nutritional status of women in general and pregnant women in particular may also improve infant growth performance during infancy, over and above any effect mediated through birth weight and birth length. The inverse relationship between maternal age and infant growth performance over the first 18 months could be related to compromised infant care. In our study setting older maternal age implies large number of children in a family, all demanding maternal care. This might dilute family resource allocation including maternal care and it might limit the amount and quality of maternal care available for the index infant. However, larger family size could also give an opportunity for shared parenting among family members, specially, after six months of age (Stewart 2007).

In this study, the mean poverty index score was significantly higher for mothers whose infants were undernourished at the age of six months or at a later age compared to mothers whose infants were not. However, the mean maternal autonomy score and mean score of availability of support to the mother did not differ between the two groups of women. In the adjusted logistic regression models none of the three composite scores were significantly related to infant stunting or underweight. At six months of infant age, contrary to our expectation, scoring higher on poverty index was positively associated with infant length

and maternal autonomy score was inversely associated with infant length and length-for-age. In LGM and MGM increased maternal autonomy was significantly associated with shorter initial length and larger total gain in length during the follow-up. In cross-sectional analysis there was a trend of inverse relationship between scoring high on poverty index and compromised infant growth after six months of age. This trend coincides with the timing that infants should be getting complementary feeding show better growth performance. As the primary caregiver, it is probable that maternal autonomy is critical for the overall wellbeing of infants (Engle, Menon et al. 1999) although it did not show significant protective effect against infant undernutrition in this study except its positive effect on a total gain in length during the follow-up. In a cohort study in Pakistan (Rahman, Iqbal et al. 2004) maternal financial autonomy was not significantly associated with nutritional status of infants after adjusting for other risk factors but higher socioeconomic status was associated with better nutritional status. In India, better nutritional status was associated with increased maternal autonomy to use household finances and increased freedom to go to the market (Doan and Bisharat 1990; Shroff, Griffiths et al. 2009). In another urban sample of under five children in India, of which 54.4% were 0-12 months old, the prevalence of underweight was higher among middle income groups relative to the lower income group (Kumar, Goel et al. 2006). Socioeconomic differentials (Ashworth, Morris et al. 1997; Baxter-Jones, Cardy et al. 1999) and the amount of social support during pregnancy (Hoffman and Hatch 1996; Feldman, Dunkel-Schetter et al. 2000) influence foetal growth, with an unfavourable postnatal environment aggravating the situation during infancy and beyond (Kalanda, van Buuren et al. 2005). A significant reduction in undernutrition among under five children in Brazil was attributed to the overall economic development in the country (Monteiro, Benicio et al. 2009). After adjusting for the effect mediated through birth weight these constructs might have only exerted a small influence in our study area and we may not have been power to detect such small effects. In other words, the overall level of poverty in the current study area, with very low variability of socio-economic status amongst participants and a high rate of non-literacy rate that might not have allowed mothers to exercise their reportedly high autonomy for better childcare practices, might explain the current non-significant findings. It is not obvious how maternal experience of obstetric complication during the delivery of index child could positively affect initial weight, initial weight-for-age and total gain in

length during the follow-up although the reverse causality might be one plausible explanation. It is possible that the reported obstetric complication could be the result of giving birth to a healthy grown fetus and this infant continue to perform well in its growth. Another plausible explanation might be the result of excessive infant care by the disadvantaged mother to guaranty the safety of the infant so that the hardships of the pregnancy period and delivery situations are all justified.

Low birth weight was a significant predictor of compromised infant growth at two, six and 12 months of age. In a more complex analysis of repeated growth measures an increase in birthweight was a significant predictor of better initial growth and smaller rates of linear change of weight-for-age and length-for-age. The current finding is in line with a previous study from Pakistan reporting a higher risk of stunting at one year of age for low birth weight infants (Saleemi, Ashraf et al. 2001). In the absence of birth weight and length, neonatal weight and length were the most important predictors of child nutritional status in Indonesia (Schmidt, Muslimatun et al. 2002). The effect of low birth weight we have observed at six and 12 months might be due to their disadvantage in utero (Strauss and Dietz 1998; Kalanda, van Buuren et al. 2005) or might have been mediated through recurrent infant illness episodes (Barros, Huttly et al. 1992; Kebede and Larson 1994; Lira, Ashworth et al. 1996; Kalanda, van Buuren et al. 2005). However, the former is more likely. Although, in the current study, initial growth inequality was maintained during infancy, birth weight did not significantly affect total gain in weight and length, and total change in the risk of underweight and stunting between two and 18 months of age.

Living in poor sanitary conditions significantly increased the risk of an infant having compromised weight at 6 and 12 months of age and compromised length at 12 months of age. However, at the age of two months the association was positive with length and marginally negative with weight. The findings from the analysis of repeated growth measures using LGM and MGM showed that the negative effect of living in poor sanitary conditions on weight operates during the whole period of infancy. However, its negative effect on infant length starts after the age of two months. Although these findings might be expected because of the differential sensitivity of weight and length to the environmental

factors including sanitary condition of the household, improving sanitation is an important area for increased intervention to reduce undernutrition (Bhutta, Ahmed et al. 2008). Other studies have not always found such an association (Vella, Tomkins et al. 1994; Mekonnen, Jones et al. 2005). Poor sanitation most probably exerts its effect on infant nutritional status by increasing the risk of infectious illnesses (Schmidt, Muslimatun et al. 2002). Therefore variation in factors affecting the prevalence of infectious disease between settings may explain the observed difference. Ethiopian mothers have been shown to understand the role of poor sanitation in child undernutrition but tend to attribute child undernutrition to their poor economic status and not something amenable to changes in their sanitation behaviours (Mekonnen, Jones et al. 2005).

In a cross-sectional analysis girls are lighter and shorter than boys, score better on weight-for-age and length-for-age, and less likely to be stunted during the first year of infancy and less likely to be underweight before one year of age. With respect to initial growth levels the findings of LGM and MGM were consistent with the results of cross-sectional analysis. Girls performed better than boys in their linear change of length-for-age but not in the rate of change of other growth outcomes. A significant nutritional disadvantage amongst boys as compared to girls is consistent with other Ethiopian studies (Central Statistical Authority 2004; Mekonnen, Jones et al. 2005; Central Statistical Agency and ORC Macro 2006) and other African studies (Vella, Tomkins et al. 1994; Espo, Kulmala et al. 2002; Ukwuani and Suchindran 2003; Wamani, Tylleskär et al. 2004; Wamani, Åstrøm et al. 2006; de Poel, Hosseinpoor et al. 2007; Wamani, Åstrøm et al. 2007), and more pronounced in the lowest socioeconomic groups (Wamani, Åstrøm et al. 2007). These results have not, however, been confirmed in a number of other studies (Vella, Tomkins et al. 1992; Kikafunda, Walker et al. 1998; Umeta, West et al. 2003; Edris 2007). The higher prevalence of anaemia among boys and in infants from poor households in a nationally representative Ethiopia study (Central Statistical Agency and ORC Macro 2006) is in keeping with the current finding. The existing vulnerability of boy babies that is seen in all cultures, and may partly have a genetic basis (Wells 2000), may explain the observation. Better initial growth performance of girls than boys and non-significant effect of gender on total gain in growth outcomes or linear growth in weight-for-age are inline with the assumption of genetic predisposition (Wells

2000) as a plausible explanation for the observed gender effect. Gender preference and differential feeding practices or neglect of boys are unlikely to be the reason, with published studies in Ethiopia tending to show that it is female infants who usually receive less food than their male counterparts (Hadley, Lindstrom et al. 2008; Koohi-Kamali 2008).

In our cross-sectional analysis rural residence significantly predicted compromised infant growth in line with other African countries (Smith, Ruel et al. 2005; Fotso 2007). The findings from MGM and LGM also showed that the growth performance of a two month old urban dwelling average infant was significantly better than its rural counterpart. However, there was no significant rural-urban difference in the rate of linear change or total change in infant growth outcomes during the follow-up. The observed urban-rural difference in growth performance might be explained by cumulative effects of series of more favorable conditions to the urban settings including socioeconomic factors (Smith, Ruel et al. 2005). A recent demographic and health survey has revealed systematic inequality in various measures between urban and rural settings including undernutrition (Central Statistical Agency and ORC Macro 2006) confirming the result of four repeated countrywide surveys in Ethiopia which have also reported a higher prevalence of undernutrition among rural populations (Central Statistical Authority 2004). Discrepancy in the source of income, exposure to the knowledge of appropriate weaning food, and purchasing power for proper weaning food, all favoring the urban setting, could explain some of the observed differences in the current study. Overall, Ethiopian mothers seem to associate undernutrition of their infants with shortage of food (Mekonnen, Jones et al. 2005). The higher prevalence of maternal undernutrition in rural areas (Central Statistical Agency and ORC Macro 2006) is another possible explanation for the current finding although the finding is independent of maternal mid-upper arm circumference and maternal height. Differential micronutrient intake in the two settings is less likely to be an explanation, although consumption of low levels of micronutrient rich food in general might explain part of the high prevalence of undernutrition in the whole birth cohort (Central Statistical Agency and ORC Macro 2006).

In cross-sectional analysis of the current study growth levels attained at two months of age were not significantly associated with the number of prior siblings aged less than 5 years.

However, having one previous child aged less than 5 years was found to be protective of being underweight at six months of age which might be explained by the child caring experience of the mother with her first child. There was a statistically non-significant but consistent association with underweight and stunting at twelve months of age showing an increased risk of undernutrition with an increased number of under five children. In unadjusted MGM but not in a fully adjusted model, having previous children aged less than 5 years had significant negative effect on the rate at which infants gain their weight and positively associated with the rate that logit of underweight increases. A regional survey within Ethiopia (Yimer 2000) and a study in Vietnam (Hien and Kam 2008) have found a direct relationship between number of under five children and undernutrition, which may be explained by the negative effect of short birth intervals (i.e. explained by more number of under five children), on quality of maternal child caring practices. Limited financial resources of the household versus an increase of demand for that resource, including good parenting practices (WHO 1989), might also explain the association

8.5 Different modelling techniques from the prospective of evaluating the effect of perinatal CMD on infant growth in Butajira

8.5.1 Relevance, flexibility and accessibility of different modelling techniques

8.5.1.1 Logistic regression and linear regression

While investigating the effect of CMD on child growth the common practice in LAMIC is to use logistic regression for binary growth outcomes (Stewart 2007; Surkan, Kawachi et al. 2008) and linear regression for continuous growth outcomes (Tomlinson, Cooper et al. 2006; Stewart, Umar et al. 2008; Black, Baqui et al. 2009). These modeling techniques are widely implemented in commercially available software packages and hence easily accessible to researchers. The minimum data required to run logistic/linear regression is a one time point

information obtained from a study participating child (eg growth measurement) and the primary guardian (eg. measurement of CMD) which can be generated from a cross-sectional surveys. The two modeling techniques (a) are covered in most introductory epidemiology and biostatistics courses and hence easily accessible to biomedical researchers, (b) can be easily used by researchers with minimum exposure training of commercially available software packages like STATA and SPSS, (c) produce parameter estimates which can be easily interpreted (i.e. odds ratio from logistic regression and expected change in mean growth performance in response to a unit change in the exposure of interest) and (d) are the only alternative methods to assess the effect of CMD on child growth if there is outcome data (i.e. growth measures) at only one time point.

Our growth outcome measures were recorded at several time points and other modeling techniques like MGM and LGM could be more appropriate (Singer and Willett 2003; Bollen and Curran 2006). However, several questions of our interest can also be answered if we use logistic/linear regression by fixing the time point for analysis (Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006). Available evidences about the negative effect of CMD on child growth in LAMIC is mainly generated using these modeling techniques (Stewart, Umar et al. 2008; Black, Baqui et al. 2009; Surkan, Kennedy et al. 2011). Hence, the two modeling techniques have some degree of relevance to our current research. However, there are some key questions that these modeling techniques are not able to answer.

Several shortcomings of the two modeling techniques can be mentioned. In any epidemiological investigation like ours missing data is unavoidable reality. For example, in the current study we were not able to record birth weight in four sub-districts resulting in a substantial number of missing data on this variable. Default settings of frequently used software packages to fit linear/logistic regression including STATA and SPSS handle missing data using list-wise deletion method which can potentially reduce sample size significantly and bias the result. However, in literature, there are several ways of handling missing data including mean substitution, MI and FIML, and the latter two methods are the recommended ones (Abraham and Russell 2004). Although MI can be used in cross-sectional analysis (i.e. in linear/logistic regression modeling) the common practice of

researchers in LAMIC while evaluating the effect of CMD on child growth is to use the default settings of the software packages (Stewart 2007). In this thesis we also followed the same tradition for comparability of results while modeling the effect of CMD on growth of infants at two, six and 12 months of age. That implies that the results from the different methods are based on different subsets of the data. In an ideal world one would repeat them on the same dataset to see what happened. To minimize the impact of large number of missing birthweight we included a three category variable (normal, low birth weight, not measured) in our cross-sectional analysis and the effect of low birth weight was in the expected direction. However, it would be more advantages if missing birth weight was handled using a more disciplined manner (e.g. using MI).

Linear/logistic regressions are cross-sectional by nature and they cannot properly evaluate the effect of CMD on growth trajectories. To overcome this shortcoming previous researchers have analyzed the effect of CMD on child growth at different ages of children and generated useful information (Rahman, Iqbal et al. 2004; Tomlinson, Cooper et al. 2006; Adewuya, Ola et al. 2008). If the interest is to evaluate the effect of CMD on change of growth between two time points (Hazarika 2010) or on the average growth performance after adjusting for some sort of clustering (Harpham, Huttly et al. 2005; Surkan, Ryan et al. 2007) it is still possible to use these modeling techniques. However, the application may be restricted to just two time points or one level of clustering.

Another shortcoming of logistic/linear regression is related to the way that the role of mediating variable is evaluated. For example it is meaningful to hypothesize mediating roles of birth weight, birth length and early infant illness episodes while investigating the effect of CMD on child growth (Rahman, Harrington et al. 2002; Patel, Rahman et al. 2004; Stewart 2007). However, it is not straight forward to address the hypothesis of mediation effect and decompose the total effect of CMD in to direct and indirect effects within the umbrella of logistic/linear regression analysis. The common practice is to include the hypothesized mediating variable in to the fully adjusted model and monitor the resulting effect on the coefficient of CMD measure (Surkan, Kawachi et al. 2008; Black, Baqui et al. 2009).

8.5.1.2 Multilevel Growth Modeling and Latent Growth Modeling

Unlike logistic/linear regression models which are appropriate to assess risk factors of growth level attained at a given age MGM and LGM are capable of simultaneously modeling growth performance of an individual overtime and growth performance across different individuals using growth measurement recorded at more than one time point per individual (Singer and Willett 2003; Bollen and Curran 2006). Having a larger number of growth measurements per individual is helpful to generate reliable estimates of the parameters of growth function within an individual. Assuming the appropriate dataset the first step in applying either LGM or MGM is to decide on the best fitting growth function of an individual overtime and the second step is to regress the estimates of the parameters of the growth function on pre-specified risk factors (Singer and Willett 2003; Bollen and Curran 2006). In both modeling techniques, it is possible to have the same or different predictors of subject specific growth parameters (e.g. predictors of initial value and rate of linear change in case of linear growth function). Although the approaches are somewhat different it is possible to give meaningful interpretation to the intercept term of growth trajectory while using both modeling techniques. For example, in the current study we centered our growth recording time at two months and the intercept was interpreted as the initial growth level.

MGM and LGM are the techniques developed from different perspectives that extend generalized linear models (GLM) (Bollen 1989; Raudenbush and Bryk 2002). Multilevel modeling extends GLMs by relaxing the assumption of independence of residuals and allowing several levels of data nesting (Raudenbush and Bryk 2002). Hence, MGM approach the analysis of growth from a nested data perspective (Bryk and Raudenbush 1987) acknowledging that growth measurements overtime are nested within an individual (Singer and Willett 2003). On the other hand, LGM is developed from the perspective of SEM (Bollen 1989) and approaches the analysis of growth through the use of multiple indicator latent factors (Meredith and Tisak 1990). Whereas in MGM time is treated as an independent variable at the lowest level (Singer and Willett 2003) (e.g. at infant level in our case), in LGM it is incorporated as specific constrained values for the factor loadings of the latent

variable that represents the slope of the growth curve (Bollen and Curran 2006). Both modeling techniques are capable of incorporating time-varying covariate of growth. In recent years the equivalence of LGM and MGM was recognized for a one level of nesting (Curran 2003; Willett and Bub 2005). The similarities of our growth parameter estimates of length-for-age and weight-for-age z-scores obtained from quadratic LGM and quadratic MGM is inline with that equivalence. The growth parameters of other outcomes were not comparable because of the difference in the fitted models (i.e. non-linear which is not polynomial in LGM and polynomial functions in MGM). The possibility of using flexibility of SEM after formulating MGM within the framework of SEM without restricting the levels of nesting was also demonstrated (Curran 2003).

However, as it is currently practiced by most applied researchers the two modeling techniques follow different procedures to arrive at a final model. In MGM the selection of alternative models are based on information criteria, likelihood ratio and statistical significance of parameter estimates (Rabe-Hesketh and Skrondal 2008). However, in LGM the first step is to evaluate the overall goodness-of-fit of the model using combination of fit indices and then to check if the estimates of the parameters are plausible and statistically significant (Bollen and Curran 2006). Hence, there is always a potential for the two modeling techniques to arrive at a final model having different number of predictor variables. For example, in this thesis we were able to adjust the effect of CMD for range of covariates regardless of their statistical significance. However, in LGM some covariates were removed from the model while looking for the best fitting conditional LGM using fit indices as the criteria. The unconditional growth models of four growth outcomes had also different functional forms under MGM and LGM.

With the assumption of missing at random the software packages designed to fit MGM and LGM treat missing data using full information maximum likelihood. In comparison to logistic/linear regression analysis these modeling techniques are more complex to understand, difficult to be used by less experienced researchers and less accessible to researchers of LAMIC. In developing countries these techniques might not be included with biostatistics or epidemiology courses of the undergraduate or postgraduate programs and that

is the case in Ethiopia. Application of LGM requires familiarity with specialized commercial software packages like AMOS (Arbuckle 1995-2006) and MPlus (Muthen and Muthen 1998-2009) limiting their accessibility to researchers working in a resource limited settings. An alternative with regard to cost might be to use open source package OpenMx which runs under the freely available open-source R package and currently in a state of active development. More accessible software packages including STATA and SPSS can be used to fit MGM.

The two modeling techniques differ in their flexibility on the timing that the repeated growth measurements are recorded. The main difference emanate from the way the data are structured for the analysis; in LGM the time point of growth measurement (e.g. two month in this thesis) is considered as a variable (i.e. wide format) while in the MGM the growth outcome (e.g. infant length in this thesis) is considered as a variable and repeated growth measures on an individual are identified by subject specific identification number (i.e. long format). MGM is not limited by unequally spaced growth measurements or unequal number of repeated growth measures across infants. However, although LGM can handle unequal number of growth measurements of different infants it requires all infants to have their growth measurements at the same time points or at least within comparable time range. For example, in our LGM we created a four month time point from all growth measures recorded between three and five months of age. However, exact time of growth measurements were used in MGM without requiring creation of any additional time point.

MGM and LGM also differ in the flexibility and richness of growth functions that can be fitted to describe growth trajectory of individual infant overtime. In MGM polynomial functions of various degrees are frequently used as candidate growth models to describe growth performance of individual infant overtime(Singer and Willett 2003). Theoretically every parameter of the growth function can be treated as random although the number could be limited by the available number of data points. In LGM it is possible to specify polynomial growth functions as well as non-linear growth function whose shape is determined by the data (Bollen 1989). For example, in this thesis we made selection of the best fitting unconditional latent growth function for an average infant after fitting

polynomial growth function with maximum degree of two and non-linear growth function which is different from polynomial functions. However, non-linear function which is different from the polynomial function was not among the candidate growth functions while working with MGM. These differences resulted in different unconditional growth functions for four growth outcomes (i.e. two unstandadized and two binary growth outcomes) under the two modeling techniques.

Another key difference of LGM and multilevel MGM is the way that mediating variables are treated within the conditional growth models. In conditional LGM all hypothesized pathways of effects are specified, appropriate regression equations are specified and systems of equations are solved simultaneously. By doing this it becomes possible to estimate direct and indirect effects of the exposure variable on the outcome variable, and direct effect of the hypothesized mediating variable on the outcome variable. For example, in this thesis, we hypothesized significant mediating role of birth weight and early infant feeding practices in the association of antenatal CMD and infant growth. Hence, direct and mediated effects of antenatal CMD and direct effects of the hypothesized mediating variables on infant growth were estimated and tested for their statistical significance. However, conditional MGM does not have this type of flexibility. One possibility was to include hypothesized mediator variable into fully adjusted conditional MGM and monitor if there is significant change in the magnitude of the effect of the target exposure variable. For example, in this thesis, this would mean monitoring the effect of antenatal CMD on initial value and rate of linear change of infant growth outcomes as the result of including one potential mediating variable (eg. delayed initiation of breast feeding or birth weight). However, this approach does not help to decompose the effect of antenatal CMD into direct and indirect effects. Another alternative approach would be to fit several MGMs in such a way that direct and indirect effects of CMD could be estimated.

8.5.2 Knowledge obtained from the findings of different modelling techniques about the predictors of infant growth in Butajira

8.5.2.1 Risk factor specific comments in line with the modelling techniques

Perinatal CMD: In logistic/linear regression analysis there was no clear picture of whether perinatal CMD has any meaningful effect on growth performance of infants except some indication of positive effect of remitted antenatal CMD on length(in cm), negative effect of postnatal CMD on length-for-age z-score, and positive effect of postnatal or chronic CMD on weight compared to the no exposure to perinatal CMD group. In MGM there was a non-significant trend showing a negative effect of perinatal CMD on initial length and positive effect on initial weight. In LGM antenatal CMD had positive effect on initial weight-for-age and length-for-age z-scores, and negative effects on total gain in weight, on the rates of linear change of weight-for-age and length-for-age z-scores. Postnatal CMD was significantly associated with compromised initial length-for-age and increased initial risk of stunting, and with better performance in these two growth measures afterwards. Persistent CMD was significantly associated with increased initial risk of stunting, small change in the risk of stunting during the follow-up, and inversely associated with the linear change of length-for-age through its positive effect on the occurrence of diarrhoea episodes.

Early infant feeding practices: Early infant feeding practices did not mediate the effect of CMD on infant growth. All modelling techniques identified receiving colostrum as a risk factor for a compromised initial length and initial length-for-age. In linear regression receiving pre-lacteal feeding was identified as a risk factor for compromised weight-for-age at the age of two month and in MGM it was identified as a risk factor for compromised initial length-for-age. Delayed initiation of breastfeeding had (a) negative effect on weight-for-age at the age of six months in linear regression, (b) negative effect on initial weight and initial weight-for-age in MGM, and (c) positive effect on linear change of length-for-age and negative effect on the change in the probit of stunting between two and 18 months in LGM

Birth weight: Birth weight did not mediate the effect of antenatal CMD on infant growth performance. In linear/logistic regression and in MGM low birthweight had significant negative effect on initial weight and less clear negative effect on initial length. However, in LGM birth weight was a significant predictor of initial growth performance (i.e. length and weight in all forms) and inversely associated with linear change of length-for-age and

weight-for-age. In linear/logistic regression the negative effect of low birth weight on both length and weight are evident at six and 12 months of age.

Infant gender: In logistic/linear regression boys perform better in weight (in kg) and length (in cm) and girls perform better in standardized growth measures during the whole year of infancy. In MGM and LGM boys perform better in initial weight (in kg) and in initial length (in cm) and girls perform better in initial values of standardized growth measures. Although there was no significant effect of gender on the rate of growth performance in MGM girls perform better in linear change of length-for-age in LGM.

Maternal height and MUAC during pregnancy: In logistic/linear regression and in LGM maternal height and MUAC during pregnancy had significant positive effect on the two month infant weight. Similarly, in MGM and in linear/logistic regression maternal MUAC during pregnancy was significantly associated with better performance of initial growth and attained growth levels starting at six months, respectively, of all growth outcomes.

Maternal Autonomy scale: Higher maternal autonomy score was (a) a risk factor for compromised infant length at the age of six months in linear regression and (b) a risk factor for compromised initial length and positively associated with the gain in length during infancy in LGM and MGM.

Obstetric complication: In all models experiencing at least one obstetric complication was significantly associated with better initial weight. The positive effect on weight which remained for the whole year of infancy in linear/logistic regression shows that infants have maintained their initial weight difference. Significant positive effect was also observed on the change of length (in cm) in LGM.

Residence: In all modelling techniques rural residence had significant negative effect on initial length and not significantly associated with initial weight. In linear/logistic regression the negative effect on length was maintained at one year of age and in MGM rural residence

had significant negative effect on the linear change of length(in cm) but not significantly associated with the change in length-for-age and logit of stunting.

Poor sanitary condition scale and poverty scale: Different modelling techniques resulted in a comparable result. All of them identified higher score on poor sanitary condition scale as having positive effect on the initial length and as being a risk factor for compromised gain in length and weight during infancy although the latter is less clear in MGM. It is also indentified as a risk factor of initial weight in MGM and a risk factor of initial level of underweight in MGM. There was an indication for inverse relationship in the results of logistic/linear regression between higher score on poverty scale and infant growth after six months of age. However, these associations were not significant in MGM and poverty scale was removed from the final LGM of four growth outcomes.

8.5.2.2 Summary

The findings about the effects of CMD on infant growth were not uniform across different modelling techniques. In logistic/linear regression there were indications of counter intuitive positive effects of CMD on infant growth, specially, on infant weight and in MGM CMD had no significant negative effect on infant growth. In LGM there were indications of (a) positive and negative effects of antenatal CMD on initial growth and on growth performance after two months of age, (b) negative and positive effects of postnatal CMD on initial length and performance of infants in length after two months of age, and (c) persistent CMD was significantly associated with an increased initial risk of stunting with smaller change of the risk during the follow-up, and inversely associated with the linear change of length-for-age through its positive effect on the occurrence of diarrhoea episodes.

After taking account of CMD similar risk factors of compromised infant growth were identified by different modelling techniques (see tables 8.1-8.4 for comparison of LGM and MGM in terms of significant predictors of initial values and rate of linear change). Clearly identified risk factors were low birthweight, being male infant, scoring higher on a poor sanitary condition scale, and reduced maternal MUAC during pregnancy. Less obvious risk

factors include sub-optimal infant feeding practices and scoring higher on poverty index. Two counter intuitive associations were also indentified: negative effect of receiving colostrum on length-for-age at two months of age and positive effect of obstetric complication on infant weight which lasted for the whole year of infancy.

Clarity of the findings to answer "which risk factor significantly affects which growth outcome and at what age?" differs among different modelling techniques, and the hierarchy in terms of preference is LGM followed by MGM and then logistic/linear regression. The findings from linear/logistic regression do not identify the effects that occurred in early infancy and then maintained over the follow-up time. Good examples from the current study include the effects of maternal MUAC during pregnancy, gender, birthweight, and obstetric complications. The difference becomes clearer in MGM and LGM because both modelling techniques allow to have different predictors for initial value (i.e. intercept) and rate of change over time (i.e. slope). LGM has an additional advantage of allowing direct inclusion of potential mediating variables into the model. Hence, in this study we were able to investigate the mediating roles of early infant feeding practices, birthweight, and infant illness while investigating the effects of CMD on child growth. We were also able to evaluate direct and indirect effect of persistent CMD through infant illness on infant growth.

Table 8.1: Significant predictors of initial length in LGM and MGM

Predictor variable	*Length*		*Length-for-age Z*		*Logit/probit of stunting*	
	LGM	MGM	LGM	MGM	LGM	MGM
Mid upper arm circumference (cm)	NS	0.117(0.042)	NS	0.047(0.017)	NS	-0.081(0.030)
Autonomy scale (0-5)	-0.193(0.060)	-0.155(0.005)	NS	-0.056(0.025)	NS	NS
Urban residence	1.140(0.392)	1.671(0.463)	0.632(0.191)	0.785(0.197)	-1.050(0.426)	-0.984(0.385)
Poor sanitary condition(0-3)	0.371(0.135)	0.350(0.007)	0.153(0.064)	0.121(0.056)		-0.236(0.104)
Female gender	-0.650(0.219)	-0.809(0.175)	0.351(0.103)	0.467(0.071)	-0.412(0.160)	-0.903(0.127)
Birth weight	1.088(0.357)	-1.748(0.421)	0.722(0.169)	-0.566(0.171)	-0.511(0.207)	0.689(0.298)
Received pre-lacteal food	NS	NS	NS	-0.377(0.165)	NS	NS
Colostrums given	-0.591(0.282)	-0.571(0.228)	-0.301(0.131)	-0.246(0.092)	NS	NS

Figures presented in the body of the table were extracted from fully adjusted models presented in chapter 6 and 7; The numbers in the body of the table are *estimates (standard error)*; MLG = multilevel growth models, LGM = latent growth models; Estimates associated with birth-weight are expected change in mean of the growth outcomes per unit change in birth-weight if they are obtained from LGM and they are mean differences of the growth outcomes between normal birth-weight and low birth-weight infants if they are obtained from MGM; For the two binary growth outcomes the estimates are logit coefficients under MGM and probit coefficients under LGM. NS = not significant

Table 8:2: Significant predictors of linear change of length in LGM and MGM

Predictor variable	*Length*		*Length-for-age Z*		*Logit/probit of stunting*	
	LGM	MGM	LGM	MGM	LGM	MGM
Age (years)	-0.083(0.026)	-0.005(0.002)	NS	NS	NS	NS
Autonomy scale (0-5)	0.224(0.084)	0.011(0.005)	NS	NS	NS	NS
Had at least one obstetric complication	0.679(0.331)	NS	NS	NS	NS	NS
Poor sanitary condition(0-3)	-0.854(0.189)	-0.069(0.012)	-0.046(0.016)	-0.025(0.005)	0.162(0.077)	0.043(0.012)
Female gender	NS	NS	0.053(0.026)	NS	NS	NS
Diarrhoeal episode in the first 2 months	NS	NS	-0.073(0.031)	NS	0.131(0.067)	NS
Birth weight	NS	NS	-0.089(0.044)	NS	NS	NS
Breast feeding delayed for 1 hours	NS	NS	0.125(0.053)	NS	-0.179(0.076)	NS

Figures presented in the body of the table were extracted from fully adjusted models presented in chapter 6 and 7; The numbers in the body of the table are *estimates (standard error)*; MLG = multilevel growth models, LGM = latent growth models; Estimates associated with birth-weight are expected change in mean of the growth outcomes per unit change in birth-weight if they are obtained from LGM and they are mean differences of the growth outcomes between normal birth-weight and low birth-weight infants if they are obtained from MGM; For the two binary growth outcomes the estimates are logit coefficients under MGM and probit coefficients under LGM. NS = not significant

Table 8.3: Significant predictors of initial weight in LGM and MGM

Predictor variable	*Weight*		*Weight-for-age Z*		*Logit/probit of underweight*	
	LGM	MGM	LGM	MGM	LGM	MGM
Height (cm)	0.011(0.004)	NS	0.016(0.006)	NS	-0.015(0.096)	NS
Mid upper arm circumference (cm)	0.034(0.012)	0.065(0.012)	NS	0.077(0.014)	-0.084(0.031)	NS
Had at least one obstetric complication	0.107(0.052)	0.148(0.052)	NS	0.174(0.063)	-0.255(0.128)	NS
Poor sanitary condition(0-3)	NS	-0.074(0.032)	NS	-0.127(0.043)	0.200(0.073)	NS
Female gender	-0.226(0.049)	-0.311(0.049)	0.342(0.073)	0.311(0.060)	-0.613(0.17)	-0.820(0.168)
Birth weight	0.592(0.075)	-0.660(0.117)	1.023(0.111)	-0.864(0.144)	-0.753(0.176)	1.750(0.338)
Breast feeding delayed for 1 hours	NS	-0.145(0.053)	NS	-0.144(0.065)	NS	NS

Figures presented in the body of the table were extracted from fully adjusted models presented in chapter 6 and 7; The numbers in the body of the table are *estimates (standard error)*; MLG = multilevel growth models, LGM = latent growth models; Estimates associated with birth-weight are expected change in mean of the growth outcomes per unit change in birth-weight if they are obtained from LGM and they are mean differences of the growth outcomes between normal birth-weight and low birth-weight infants if they are obtained from MGM; For the two binary growth outcomes the estimates are logit coefficients under MGM and probit coefficients under LGM. NS = not significant

Table 8:4: Significant predictors of linear change of weight in LGM and MGM

Predictor variable	*Weight*		*Weight-for-age Z*	
	LGM	MGM	LGM	MGM
Age (years)	-0.027(0.008)	-0.002(0.001)	-0.004(0.002)	-0.002(0.001)
Poor sanitary condition(0-3)	-0.158(0.054)	-0.008(0.004)	-0.031(0.011)	NS
Birth weight	NS	NS	-0.111(0.028)	NS

Figures presented in the body of the table were extracted from fully adjusted models presented in chapter 6 and 7; The numbers in the body of the table are *estimates (standard error)*; MLG = multilevel growth models, LGM = latent growth models; Estimates associated with birth-weight are expected change in mean of the growth outcomes per unit change in birth-weight if they are obtained from LGM and they are mean differences of the growth outcomes between normal birth-weight and low birth-weight infants if they are obtained from MGM; For the two binary growth outcomes the estimates are logit coefficients under MGM and probit coefficients under LGM. NS = not significant

8.6. Contribution of the current research work

8.6.1 Contribution to the existing knowledge

- While antenatal CMD had direct positive effect on initial growth and direct negative effect on the rate with which infant grow overtime, postnatal and persistent CMD were significant risk factors of compromised infant length. The later effect on the total gain in length during the follow-up was partly mediated through diarrhoeal episodes.
- Neither early infant feeding practices nor birthweight were significant mediators of the effect of CMD on infant growth. However, there was some degress of association of compromised infant length with receiving clostrum and pre-lacteal feeding, and delayed initiation of breast.
- Boys perform better in initial weight (in kg) and length (in cm), and girls perform better in initial values of standardized growth measures, and these differences were maintained during the first 18 months of life.
- Low birthweight and reduced MUAC during pregnancy were significant predictors of compromised initial growth and these effects were maintened over the first 18 months of infancy.
- Rural residence was significant predictor of better initial length. However, there was no significant urban /rural differential in initial weight and on how infant grow overtime
- Higher maternal autonomy score was a risk factor for compromised initial length and it was positively associated with the total gain in length during infancy
- Maternal experience of at least one obstetric complication was significantly associated with higher initial weight, and the statistical significance of the weight difference was mainted during the first year of infancy.
- Higher score on poor sanitary condition scale was significantly assocated with better initial length and compromised initial weight. However, it was inversely associated with the change in weight and length during infancy.

- There was no clear association of scoring higher on poverty scale and infant growth although there was some indication of inverse relationship after six months of age.
- Of the three modelling techniques, LGM has more flexibility and produces easly interprateable results. In a situations where LGM is not accessable MLG is the preferred alternative over the cross-sectional modelling techniques (i.e. logistic/linear regression)

8.6.2 Knowledge transfer to the low income country setting in terms of modelling capability

While developing this thesis the PhD candidate has got an opportunity to read and apply statistical methodologies which are currently used in developed countries to model repeated measures. These methodologies are not yet extensively used in LAMIC to model the interrelationships of maternal mental health and child physical health. The most likely reason why these methodologies are under utilized in LAMIC is lack of exposure to these methods in the universities, shortage of expertise in the field and limitation of resources to access appropriate software packages. The candidate has now acquired knowledge of the theoretical background to these methods and considerable practical experience in collecting suitable data in the field and in applying the statistical methods. He is also based within the biomedical research Institute of Addis Ababa University and his strong link with the Institute of Psychiatry will be maintained. Department of Psychiatry of the Addis Ababa University has planned to start a new PhD program in mental Health Epidemiology in October 2011. These conditions will give the PhD candidate an opportunity to introduce MGM and LGM into the postgraduate teaching and research programs of College of Health Sciences, Addis Ababa University. However, there will be challenges to the candidate while trying to introduce the application of these statistical methodologies within Addis Ababa University and beyond. Hence, he should be able to extend his external collaboration to get assistance in case of difficulties related to the theory and application of these methodologies.

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